**UNIVERSITÀ DEGLI STUDI
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Local and National perspectives on Poverty, Fertility, and Social Mobility

Candidato:

Andrea Barigazzi

Relatore (Tutor):

Prof. Massimo Baldini

Correlatore (Co-Tutor):

Dott. Giovanni Gallo

Coordinatrice del Corso:

Prof.ssa Ylenia Curzi

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Abstract (lingua italiana)

La presente tesi di dottorato raccoglie tre contributi nell'ambito dell'economia pubblica, cercando di allargare la letteratura esistente di tre tematiche cruciali: povertà, mobilità sociale e fecondità.

Il primo contributo studia le modifiche nelle richieste di benefici sociali durante la pandemia da Covid-19 in Italia. L'indagine valuta se la pandemia abbia ridotto i costi di transazione per l'accesso ai trasferimenti sociali, ampliandone la diffusione nella popolazione, anche considerando le tendenze recessive. I risultati, basati su dati dettagliati a livello regionale, rivelano una correlazione significativa e positiva tra la diffusione dei casi di COVID-19 e la distribuzione dei benefici sociali, mettendo in luce l'impatto delle misure di lockdown sul benessere economico delle famiglie e di conseguenza sui costi legati alle richieste di trasferimenti sociali.

Il secondo contributo si concentra sulla mobilità sociale, introducendo una fonte innovativa di dati: i giornali. Ipotizzando che gli individui frequentemente menzionati nei giornali abbiano rilevanza sociale, l'analisi si concentra sui cognomi nei giornali della provincia di Modena, Italia, dal 1921 al 2011. La rappresentazione relativa di tali cognomi rispetto alla loro presenza nella popolazione generale rivela una sovra-rappresentazione di gruppi elitari nei quotidiani, che tende a perdurare nel tempo. Questo suggerisce una trasmissione intergenerazionale dello status sociale. Tale analisi contribuisce a identificare uno strumento alternativo nello studio della mobilità, specialmente in assenza di dati reddituali.

Il terzo contributo esamina la relazione tra fecondità e politiche famigliari nei paesi dell'Unione Europea. Basandosi sui dati EU-SILC del periodo 2004-2020, la ricerca indaga come cambiamenti rilevanti nella fornitura di trasferimenti familiari influenzino la probabilità di avere un figlio nel breve termine. Basandosi sull'ipotesi dell'investimento sociale e su una generale espansione delle politiche familiari nell'UE dal 2004, esploriamo se e come il maggior sostegno alle famiglie contribuisca agli eventi nascita all'interno delle famiglie. L'analisi valuta le reazioni asimmetriche nella fertilità ai cambiamenti nei benefici sociali legati alla famiglia, guardando specificamente ai cambiamenti nei loro livelli di generosità e universalismo di anno in anno. I risultati indicano che l'aumento della generosità dei benefici monetari è positivamente correlato con un aumento della probabilità di avere un figlio, mentre l'opposto si verifica in caso di riduzione dell'universalismo dei trasferimenti. I risultati dell'analisi dell'eterogeneità rivelano anche che la reazione ai cambiamenti nella spesa pubblica non è la stessa in tutta la popolazione, ma varia secondo le caratteristiche delle madri.

Abstract (lingua inglese)

This PhD thesis comprises three contributions in the field of public economics, aiming to expand existing literature on three crucial themes: poverty, social mobility, and fertility.

The first contribution examines changes in claim for social benefits during the Covid-19 pandemic in Italy. The study assesses whether the pandemic has reduced transaction costs for accessing social transfers, thereby increasing their dissemination in the population, even considering recessionary trends. Results, based on regional data, reveal a significant and positive correlation between the spread of Covid-19 cases and the distribution of social benefits, highlighting the impact of lockdown measures on the economic well-being of families and consequently on the costs associated with social transfer claim.

The second paper focuses on social mobility, introducing an innovative data source: newspapers. Assuming that individuals frequently mentioned in newspapers have social relevance, the analysis centers on surnames in newspapers from the province of Modena, Italy, spanning from 1921 to 2011. The relative representation of these surnames compared to their presence in the general population reveals an overrepresentation of elite groups in newspapers, which tends to persist over time. This suggests intergenerational transmission of social status. This analysis contributes to identifying an alternative tool in the study of mobility, especially in the absence of income data.

The third contribution examines the relationship between fertility and social policies across countries within the European Union. Based on 2004-2020 EU-SILC data, the research investigates how relevant changes in family/children allowances influence the likelihood of new births in the short term. Based on the social investment hypothesis and a general expansion of family policies in the EU since 2004, we investigate if and how increased family support contributes to birth events within families. The analysis assesses asymmetric reactions in fertility to changes in family-related social benefits, specifically looking at changes in their generosity and universalism levels from one year to another. Findings indicate that enhancing the generosity of cash benefits is positively correlated with an increase in the likelihood of having a child, while the opposite occurs in case of reduction of the transfers universalism. Results of the heterogeneity analysis also reveal that the reaction to changes in public spending is not the same across the population but differs according to characteristics of mothers.

Chapter 1:
How Pandemic Shock Affects Claim for Minimum Income Measures*

Abstract

This study aims to understand how the demand for the minimum income measure in Italy has changed in response to the greatest crisis of recent years, namely the pandemic shock. In particular, we hypothesize that the pandemic has reduced the transaction costs associated with claiming social transfers, increasing their spread across the population even when controlling for recent recessive trends. We focus on Italy as an interesting case study because it was the first Western country to be strongly affected by the Covid-19 pandemic and the last EU country to introduce a national minimum income scheme (the Citizenship Income or RDC). Based on a rich set of NUTS-3 regional-level statistical data, the results show a significant and positive correlation between the spread of RDC recipients and Covid-19 infections, especially during the first phase of the pandemic. This evidence confirms that lockdown measures have strongly influenced the economic well-being of households. In addition to expanding the number of families in difficulty, this may have indirectly led to a reduction in the transaction costs associated with applying for the RDC benefit. The main results hold when considering relevant demographic and socioeconomic variables that directly influence the RDC claim.

Keywords: Minimum income schemes; transaction costs; Covid-19; social transfers; NUTS-3 regions.

JEL classification codes: I18; I31; I38.

* This chapter is co-authored with Giovanni Gallo (University of Modena and Reggio Emilia).

1. Introduction

Social transfers, and minimum income schemes in particular, are key tools to support people's income and protect their living standards. However, the incisiveness of such policies may be undermined if eligible recipients do not claim the benefits they are entitled to. Several studies have provided evidence of non-take-up of social policies in developed countries (Hernanz et al., 2004; Campbell et al., 2005; Figlio et al., 2015). The reasons behind insufficiently high take-up rates are multiple and they can be enclosed in the assumption that the expected benefits are too low compared to the transaction costs of claiming social assistance (Riphahn, 2001). In the literature, bureaucratic and administrative barriers are generally cited among the determinants that significantly affect the reduction in social policies take-up (Van Oorschot, 1991; Daigneault and Macé, 2020). Similarly, the expected amount and duration of benefits are major factors related to social policies take-up (Bruckmeier and Wiemers, 2012; Arrighi et al., 2015). Other works also correlate non-take-up with information and awareness among potentially eligible individuals (Matsaganis et al., 2010; Bhargava and Manoli, 2015). Not least, social stigma is indicated as an important factor leading to increased rates of non-take-up (Moffitt, 1983; Hancock et al., 2004; Baumberg, 2016).

Social transfers and minimum income schemes assume additional value during times of crisis. In fact, shocks can jeopardize the economic well-being of households, leading to economic instability, reduced purchasing power, increased uncertainty, poverty, and unemployment rate. The claim for social support is consequently expected to increase in the aftermath of a shock, independently from its kind. Several scholars have related these two aspects, studying the impact on the demand for social benefits caused by different shocks such as wars, economic crises, natural events, sudden changes in the market, or health crises. In Argentina, after the outbreak of the 2001 severe economic crisis, the government introduced the Plan Jefes, thus an income support measure for all households with workers who had lost their main source of income (Galasso and Ravallion, 2004). Between 2007 and 2009, the US public spending increased by 14.2% due to the Great Recession. Three quarters of the increase was due to the increase in cash transfers, of which three quarters, in turn, were social transfers (Oh and Reis, 2012). despite the adverse conditions related to the crisis and the severe budget austerity, the Spanish government strongly defended its minimum income scheme (established just before 2008) during the recession. Similarly, Greece and Italy, the only two EU countries without a national minimum income scheme in 2016, decided to introduce this measure in their welfare systems as a response to the recession effects in 2017 and 2018 respectively (Ziomas et al., 2017; Jessoula and Natili, 2020). Health shocks are among the most complex crises to deal with, and the Covid-19 pandemic is certainly the most impactful shock of this kind in recent years. As an example of the severity of the effects of the pandemic on national economies and labor markets, Gallo and Raitano (2023) highlight the sudden deterioration of the Italian macroeconomic situation after the arrival of Covid-19. The pandemic also elicited an inevitable and immediate response from policymakers, who supported household incomes by introducing emergency benefits or improving the existing measures (Anderson et al., 2020; Gentilini et al., 2021), temporarily putting aside the usual concerns about the trade-off between the generosity of social transfers and fiscal sustainability.

Instability resulting from shocks may lead in parallel to a reduction in transaction costs associated with the claim for social transfers. A lower non-take-up rate may depend on a stigma reduction, following the assumption that it is socially more acceptable to apply for welfare assistance when a larger percentage of the population is in economic hardship (Gustafsson, 1984; Gustafsson, 2002). At the same time, the expectation of receiving larger amounts of social assistance and for longer periods during a crisis may lead to a greater propensity to claim for social policies. Administrative procedures could also be simplified during times of economic uncertainty, as well as bureaucratic constraints could be alleviated in turn.

This paper aims to understand how the claiming of social benefits changed in response to the pandemic shock, even when controlling for its recessive impact across the national territory. To do this, we focus on Italy as an interesting case study because it was the first non-Asian country to face the rapid and widespread spread of Covid-19, the first Western country to introduce heavy restrictions on mobility and personal freedom, and the first EU country to close all activities not considered as essential (Capano, 2020; Remuzzi and Remuzzi, 2020). Moreover, Italy was the latest EU country that introduced a national minimum income scheme, i.e. the so-called Reddito di Cittadinanza or RDC (Raitano et al., 2021). Thanks to its benefit generosity, it has represented the main public policy contrasting poverty and social exclusion in Italy. For this reason, among the different cash social transfers existing in the Italian welfare system, we decide to focus on RDC in this analysis.

This study explores how the RDC claiming changed during the different stages of the pandemic in Italy from February 2020 to December 2021, as well as across the country. The latter aspect appears of great interest in the proposed analysis because the spread of Covid-19 in Italy has been quite heterogeneous at the territorial level. Based on the estimate of linear panel-data models, the econometric analysis relies on monthly data aggregated at the NUTS-3 level on the RDC receipt (e.g. number of recipient households, average benefit amount) and the Covid-19 pandemic spread (e.g. number of contagious, deaths due to the coronavirus).

To the best of our knowledge, the contribution of the paper to the economic literature on the topic is twofold. First, recent literature analyzed the state fiscal response to the pandemic shock (among others see Baptista et al., 2021), but still neglects how the claiming of social benefits has changed because of it. To do that, as the pandemic is undoubtedly a regional crisis spatially uneven in its impacts (Bailey et al., 2020; Bailey et al., 2021; Bonacini et al., 2021), we adopt a sub-regional perspective. This is particularly important in Italy, where the healthcare management is regulated on a regional basis (Mauro and Giancotti, 2021; Costa-Font and Turati, 2018). Second, we further explore the heterogeneity existing in the national territory also considering relevant socio-economic and demographic characteristics of the population at NUTS-3 regional level. These factors may indeed affect the claiming of social benefits, with some social groups more inclined to claim and others instead more reluctant to do the same (Sohrab, 1994; Currie and Grogger, 2002; Grogger and Michalopoulos, 2003). These discrepancies across the population may have several reasons, such as language difficulties, stigma, inadequate information, low program awareness, or a greater tendency to procrastinate (Lamont et al., 2014; Bruckmeier and Wiemers, 2017).

Within the context of previously discussed definitions of transaction cost, such as administrative barriers, the duration and expected amount of benefits, informational awareness among potential beneficiaries, and social stigma, it is critical to highlight that in the early months of the pandemic, the conditionalities associated with active job search and meetings with social services were temporarily suspended. Starting from April 2020, moreover, the methods for submitting applications for Citizenship Income (RDC) were expanded and more intensely promoted, allowing for submissions through the online portal of the National Institute of Social Security (INPS), whereas previously it was only possible through the digital identity (SPID, which was not widely spread at the beginning of 2020) or in person at tax assistance centers or at patronage institutes. Furthermore, the condition of economic uncertainty and the worsening financial difficulties for the entire population may have altered the perception of the social stigma associated with requesting economic support at such a delicate moment. Therefore, what we assume in this analysis is that the economic need engendered by the negative effects of the pandemic on the labor market is likely to have reduced the perceived stigma related to the RDC claim. At the same time, we expect that the suspension of the measure conditionality (i.e. the mandatory active research of an occupation) and interviews with both social services and employment centers have also decreased the fear of controls among households. As a consequence of these elements, transaction costs related to the RDC claim significantly – even if temporarily – reduced during the times of pandemic, potentially leading to a greater number of households in economic difficulties applying for the social benefit.

Our findings indicate a significant and positive relationship between RDC recipients and the trends of Covid-19 cases, suggesting that the number of RDC beneficiaries increased during periods of lockdown, which were particularly pronounced during the first wave of contagions. This evidence confirms that the massive mobility restrictions implemented by the Italian government to counter the spread of the virus have strongly impacted the economic well-being of households. This, in addition to expanding the number of families in difficulty, may have led - as a collateral effect – to a reduction in the transaction costs associated with applying for the RDC benefit.

The rest of the paper is organized as follows. Section 2 describes the datasets used and the Italian framework on the evolution of the pandemic and the RDC receipt. Sections 3 and 4 present the econometric method and results. The last section concludes and discusses policy implications arising from this study.

2. Data and the Italian framework

The analysis relies on a dataset merging, for each of the 107 Italian provinces (i.e. NUTS-3 level), aggregated statistics on the spread of Covid-19 contagions and the RDC receipt. The first ones are provided by the Italian Civil Protection Department[1] and contains information on the daily trend of positive cases and deaths from the 24th of February 2020 onwards. The second archive of aggregated statistics, named '*Osservatorio sul Reddito e Pensione di Cittadinanza*', is instead provided by the Italian National Social Security Institute (INPS) and collects several information on RDC since August 2019,[2] providing at provincial level the monthly trend of the number of households being RDC recipients and the benefit amount received on average by the same. Once merged the two datasets, as we are mainly interested on how the RDC claiming changed during the pandemic, our final sample of provincial-level observations only focuses on the period from January 2020 to December 2021.

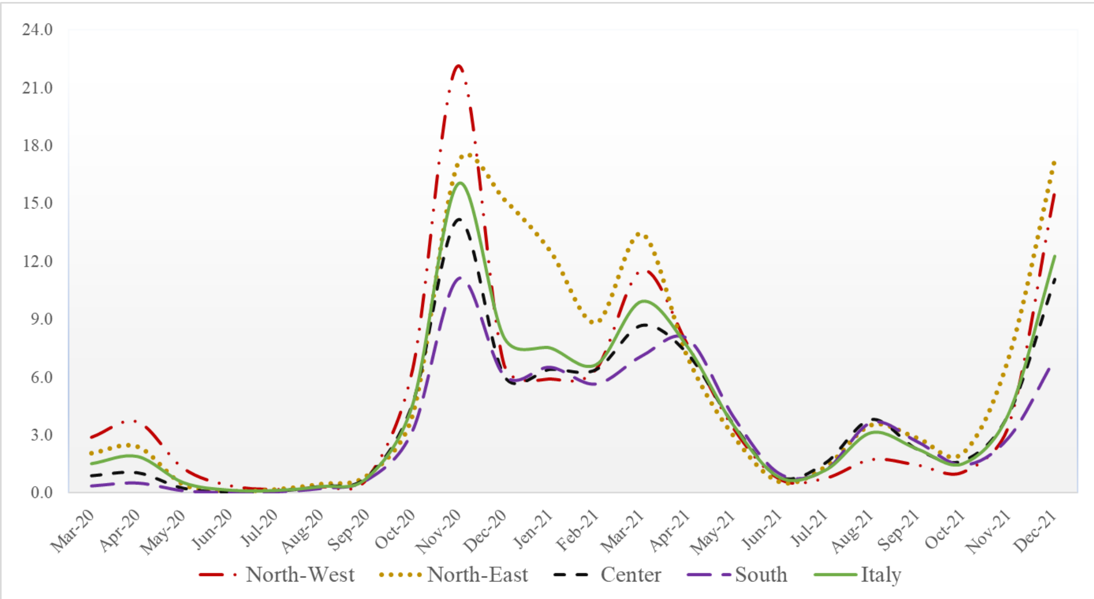
Socio-economic and demographic characteristics of the provincial population differ across the national territory. These factors are expected to claim and take-up for means-tested social benefits. This is especially true when dealing with a country like Italy, which is marked by strong heterogeneity among provinces in terms of demographic and economic characteristics (see, among others Gallo and Pagliacci, 2020). To further explore this heterogeneity in our analysis, the final dataset is enriched by a number of provincial-level statistics on relevant demographic and socio-economic characteristics of local populations. A more detailed description of variables used can be found in Appendix (Table A1).

[1] Civil Protection Department. Repository of Covid-19 outbreak data for Italy. https://github.com/pcm-dpc/Covid-19. Accessed on February 11, 2022.
[2] Although the RDC have been introduced in March 2019 (first cash payments since April 2019), the INPS provides aggregated statistics on this measure since August 2019.

2.1. The evolution of pandemic in Italy

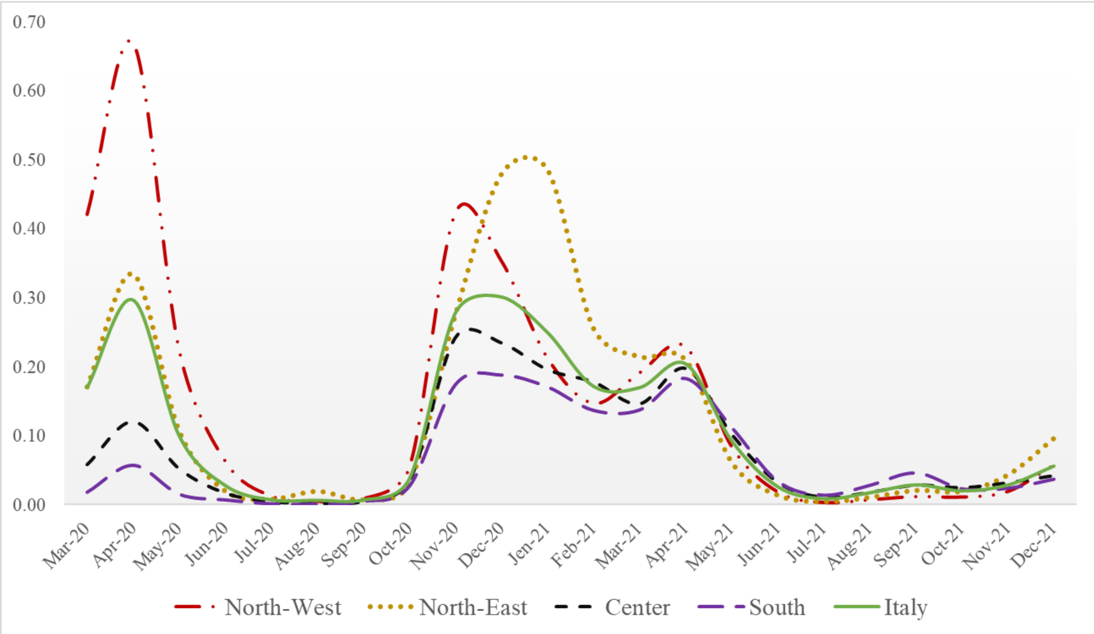
Figure 1 shows the trend of Covid-19 contagions in Italy, by macro-region (i.e. north-west, north-east, center, and south) and as a whole, between March 2020 and December 2021. As we observed different phases of coronavirus spread (and different national and local government strategies in terms of contact tracing and restrictive measures), to provide a more truthful measure of the impact of pandemic shock on local population, Figure 2 also shows the trend of deaths due to Covid-19 during the same reference period.

Figure 1. Trend of Covid-19 cases per 1,000 inhabitants between March 2020 and December 2021



Notes: The number of cases reported is the one collected on the 28th day of the month. Source: Elaborations of the authors on Civil Protection Department data (2021).

Figure 2. Trend of deaths due to Covid-19 per 1,000 inhabitants between March 2020 and December 2021



Notes: The number of deaths reported is the one collected on the 28th day of the month. Source: Elaborations of the authors on Civil Protection Department data (2021).

The first period, which includes roughly the months from March to October 2020, was initially characterized by a dramatic and unexpected influx of deaths. The sum of infections, on the other hand, when compared to the subsequent waves, seems significantly lower. This is due to the poor testing capacity, the limited availability of swabs, and the inevitable unpreparedness the local authorities faced in the early stage of the pandemic outbreak. Statistics on Covid-related deaths then return the real magnitude of the first wave of contagious. The seriousness of the situation is also confirmed by the actions taken to counter the virus spread (e.g. closure of schools and universities, ban of unnecessary travels, closure of non-essential activities).[3] These measures remained active until May 3, 2020.

The second part of this first period, which can be defined as "living with the virus", reported the gradual relaxation of restrictive measures. The situation remained stable until the beginning of November 2020, when the second wave of Covid contagious forced the Italian government to new restrictions. This period, which lasted until late spring 2021, can be considered as the second phase of pandemic. Figures 1 and 2 clearly show the exponential increase in cases and deaths since November 2020. The difference in the number of positive cases compared to the first wave is due, as mentioned, to the different tracing capacity and strategy, but also to the advent of new, more

[3] To better understand the impact of the first wave on national economic wellbeing, Figure A1 presents the time trend of the quarterly GDP amount between 2019 and 2021. The first and second quarters of 2020 mark a contraction of 6.4 percent and 18.5 percent from the same quarters of the previous year. While at the end of 2021, GDP returns to the levels of the end of 2019, before the pandemic outbreak.

contagious variants of Covid-19. Another difference with the first wave of contagious regards the kind of restrictive measures adopted. In fact, during this phase of virus expansion, instead of introducing the same measures in the whole national territory, the Italian government established a containment system where the tightening of restrictive measures was based on a set of indicators at regional level.[4] For this reason, the freeze on economic activities was in this phase more moderate than in the first one. From summer 2021 onwards, also thanks to the advent of Covid-19 vaccines and the massive vaccination of the Italian population, a stabilization of the situation and a gradual return to normality were outlined despite the arising of new Covid-19 variants.

The pandemic is undoubtedly a regional crisis, spatially uneven in its impacts. The North-West is the most affected area in the first phase of virus spread, followed by the North-East. Other areas are instead less affected, particularly the South. This is particularly clear when controlling for the number of deaths. In this case, North-West values are double compared to those in the North-East, six times higher than the Center ones, and ten times higher than the South ones. In contrast, the second wave affects the Italian macro-regions more evenly, despite the virus spread is still slightly greater in the North of Italy.

A deepening on the trend of Covid-19 at provincial level highlight the importance of studying the subregional heterogeneity. Looking at the first wave of Covid-19 cases, for instance, a subsample of the most affected Italian provinces at that stage (i.e. Lodi, Mantua, Reggio Emilia, Piacenza, Verona, and Turin) report remarkable disparities (Figure A2). The provinces of Lodi and Piacenza show similar trends in positive cases to each other despite belonging to different NUTS-1 regions (North-West and North-East respectively). Some differences between provinces also arise in terms of decrease rates: the province of Mantua, for example, reported a much faster decline in the number of Covid-19 cases and deaths with respect to the others.

2.2. The RDC receipt: characteristics, territorial distribution and recent trend

The RDC was introduced in Italy by Law No. 26/2019. Households began to apply for the measure from March 2019, and the first transfers date back to the following month. The transfer paid to households has gradually increased over time, from a monthly average of about 525 euros in 2019, to 565 in 2020, and over 580 in 2021. To be eligible for the measure, legal age of 18 and Italian or

[4] This containment system at regional level distinguished white, yellow, orange and red zones according to the seriousness of the pandemic.

EU citizenship are required. Citizens of other countries can also apply, but only if they have been resident in Italy for at least 10 years, the last two of which have been continuous.

The economic requirements are fairly stringent and require considerable administrative effort, both in terms of proof of means and documentation to be submitted. In addition to other minority requirements (mainly related to recent car/motor vehicle purchase and ship ownership), RDC has four distinct economic eligibility requirements. Specifically, the household must possess: i) ISEE (i.e. the *Equivalent Economic Situation Indicator*)[5] value of less than 9,360 €, that imposes a double administrative procedure (first for ISEE and then for RDC); ii) equivalent household income value of less than 6,000 € (9,360 € if households reside in rented houses); iii) value of movable assets not exceeding 6,000 € for a person living alone, increased according to the number of household members (up to 10,000 €); iv) value of real estate assets, other than the first house, not exceeding 30,000 €. Other requirements affect the transaction costs associated with claiming for the RDC, including willingness to tax and administrative audits, as well as declaration of immediate availability for work and adherence to an individualized job placement pathway.

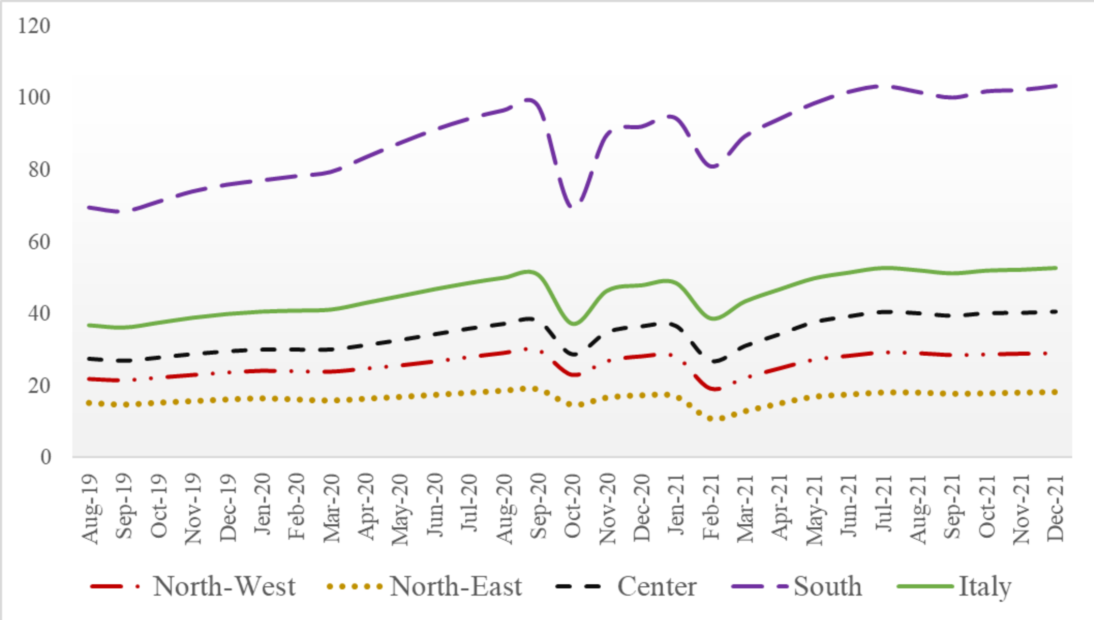
It is important to highlight that, due to the pandemic, the Decree Law of March 17, 2020 suspends for two months the conditionalities pertaining to active job search and interviews with social services. This suspension is subsequently extended for another two months by the Decree Law of May 19, 2020. The requirements are reinstated starting from mid-July 2020. Moreover, starting from April 2020, the methods of submitting applications for the RDC are expanded, which can take place online also through the website of the National Institute of Social Security (INPS). Previously, the methods of submitting applications referred to the options: online, with the condition of having the digital identity (SPID); at tax assistance centers; at patronage institutes.

Figure 3 shows that the number of RDC recipient households reported two important drops since August 2019: in October 2020 and February 2021. Both reductions are due to administrative reasons. The maximum length of the RDC receipt is 18 months, but it can be claimed again after one month break. As many households in economic need started to receive the RDC from April 2019, several of them saw the receipt expired in October 2020 (claiming again the benefit since November 2020). As for the February 2021 drop, it is instead due to the fact that a share of households often has issues in renewing documents for the annual means-test in time.[6]

[5] The ISEE is a complex indicator combining household income and wealth. It consists of the sum of the household income and 20% of the household wealth (in terms of both financial assets and property) divided by an ad hoc equivalence scale. The ISEE equivalence scale is equal to the number of household members raised to the power 0.65.

[6] The same phenomenon is not clearly visible in February 2020 because of the temporary freezing of administrative procedures due to the pandemic, but we verified that in February 2022 (which is not included in the analysis reference

Figure 3. Trends of RDC recipients per 1,000 inhabitants between August 2019 and December 2021



Source: Elaborations of the authors on INPS statistics (2021).

Despite the two anomalous drops, some interesting aspects can be highlighted. First, the number of RDC recipients' households was in December 2020 already similar to the September 2020 one, suggesting that most of those who completed the first tranche of RDC benefit claimed for a renewal in the very short run. Second, the RDC incidence varies on a geographical basis and it found to be higher in the South, where we observe more than 100 RDC recipients per 1,000 households (i.e. 10 percent of the total households) during most of 2021. Third, considering the month of December as yearly reference point, the number of RDC recipients increases much more in 2020 (20%) than in 2021 (10%). This is largely expected as the restrictive measures have been more severe and long in 2020 at national level, negatively affecting the labour market and the economy in general. Nonetheless, this relationship does not appear equally clear in some areas of the country. In fact, despite all Italian macro-regions report increasing trends in the number of RDC recipients' households after the pandemic (Table A2), the North-East present limited increases (+7% in 2020 and +5% in 2021) while being one of the areas with the greatest number of Covid-19 cases and deaths.

2.3. Investigating heterogeneities at local level

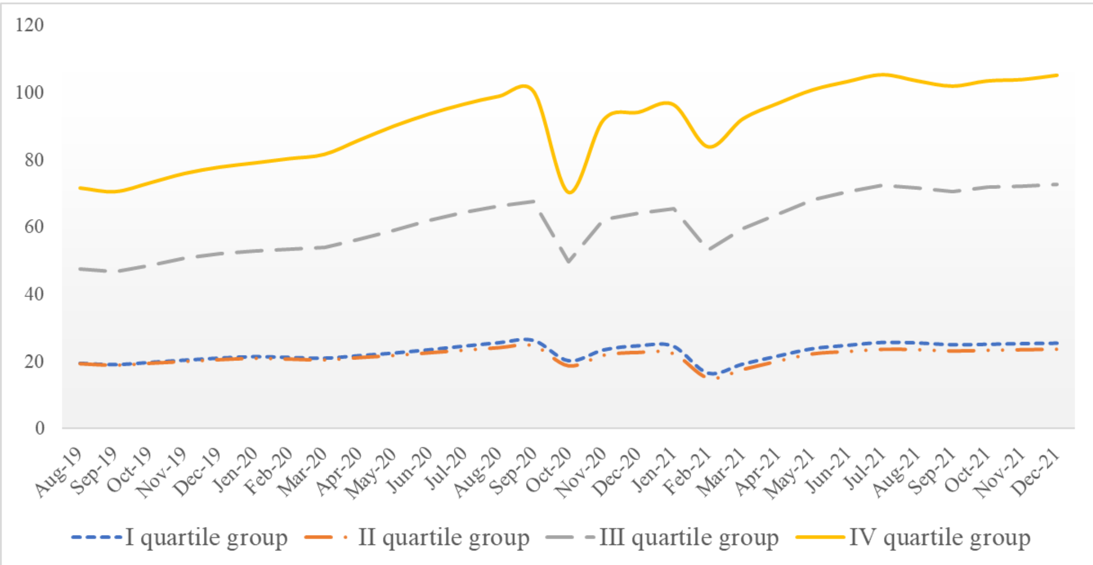
Before moving to the econometric part of the analysis, we provide here a further descriptive evidence on the relationship between RDC and Covid-19 looking at the territorial heterogeneity by income

period). From January 2022 to February 2022, we actually observe a sharp decline in the number of RDC recipients (−18% at national level) as well.

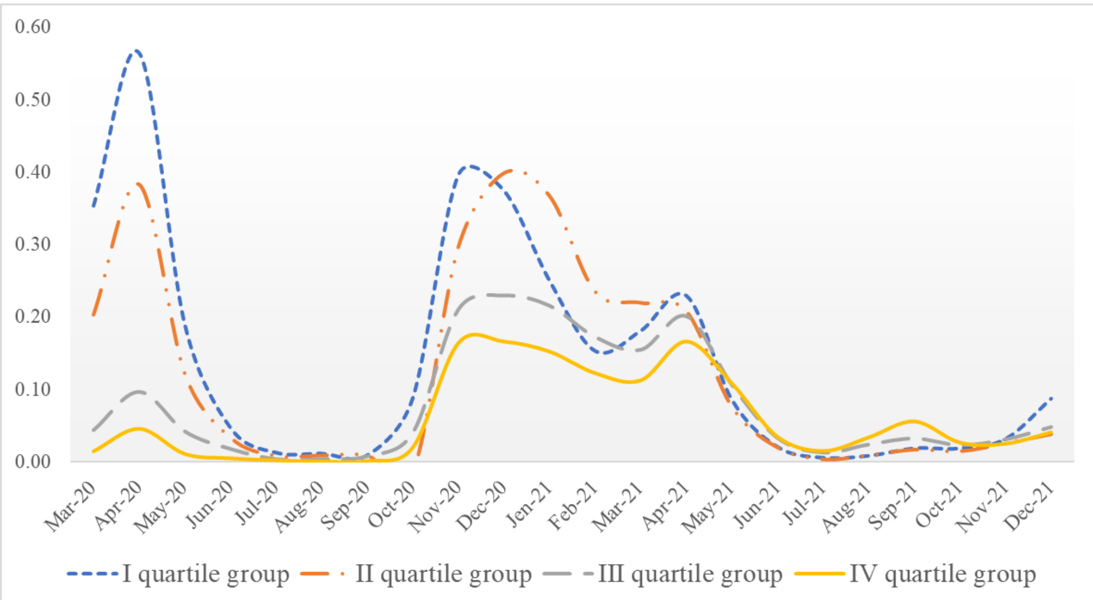
poverty. Given the unavailability of alternative poverty indicators at such regional level, this relevant dimension is here measured through the share of taxpayers declaring a taxable income lower than 10,000 € on the total number of taxpayers (Italian Ministry of Economics and Finance, 2018). In order to explore the territorial disparity in terms of poverty, Italian provinces are also divided into four groups, starting from those with the lowest poverty index (first quartile) to those with the highest poverty index (fourth quartile). Then, the trend of the territorial spread of RDC deaths due to Covid-19 is studied by poverty quartile group (panel A and panel B of Figure 4 respectively).

Figure 4. RDC recipients per 1,000 households (panel A) and deaths due to Covid-19 per 1,000 inhabitants (panel B) in Italian provinces divided by poverty index

Panel A



Panel B



Source: Elaborations of the authors on INPS data (2021), Civil Protection Department data (2021), and Minister of Economics and Finance data (2018).

Expectantly, panel A of Figure 4 shows that RDC is more spread in the poorest areas of the country (i.e. third and fourth quartiles) from the very first moment of its introduction. Restating remarks on variation in recipients in the seven months before and after the pandemic outbreak, again a greater increase in recipients in the period after the advent of Covid-19 pandemic is observed. This increase is in percentage terms very similar among the four quartiles. It is, however, larger for the poorest areas when looking at a longer time frame (March 2020 - December 2021; 21% for first quartile, 15% for second quartile, 35% for third quartile, and 28% for fourth quartile), as if to indicate a shock that initially affected both poor and rich areas, but in the medium term inflicted the worst consequences for the already poorest areas of the country. Moreover, this evidence does not seem to be related to the virus spread, as the wealthiest areas are also the territories most afflicted in terms of victims of the pandemic (panel B of Figure 4). The same results hold when replacing the adopted poverty index with the unemployment rate as indicator of territorial economic distress (Figure A3), or when referring to a demographic indicator of territorial vulnerability, thus the dependency ratio (Figure A4).

3. Econometric methods

Our basic assumption is that pandemic shocks affect the claim of social transfers. Therefore, it is expected that the evolution of pandemic trends lead to changes in the applications for the RDC benefit.

The econometric strategy is composed of two sequential parts. In the first one, we analyze the incidence of RDC recipients on provincial population in relation to the spread of Covid-19 infections and Covid-19 deaths, using fixed-effects panel data models. Regressions are distinguished by period, following the definition of different phases of virus expansion outlined in Section 3: (i) March 2020 - Sept 2020; (ii) Nov 2020 - Jan 2021; (iii) Mar 2021 – Jul 2021; (iv) Aug 2021 - Dec 2021. The months of October 2020 and February 2021 are excluded for the reasons highlighted in Section 3 regarding the decline in RDC recipients due to administrative reasons. In the second stage of the econometric analysis, to further explore heterogeneity among Italian provinces, we switch to random-effects models to account for a number of relevant (but time-invariant) demographic and socioeconomic covariates.

For the first part, we consider the following baseline panel data model specification:

$$\mathrm{Y}_{it} = \beta_0 + \beta_1 X_{i(t-1)} + \beta_2 X_{i(t-1)}^2 + \beta_3 X_{i(t-2)} + \beta_4 X_{i(t-2)}^2 + \alpha_i + \varepsilon_{it} \qquad [1]$$

Where Y_{it} is the number of RDC recipients per 1,000 households in province *i* at time *t*, $X_{i(t-1)}$ is the number of Covid-19 cases per 1,000 inhabitants in province i at time $(t-1)$, and $X_{i(t-2)}$ is the number of Covid-19 cases per 1,000 inhabitants in province *i* at time $(t-2)$. Time is considered at period $(t-1)$ and at $(t-2)$ because it is assumed that changes in the curve of infections affect the RDC claiming with a time lag due to possible administrative delays, periods of adjustment, or slowdowns in application procedures. We decide to adopt a quadratic polynomial form to test whether the relationship between Covid-19 positive cases and the number of RDC recipients has a nonlinear shape (e.g. it grows at an increasing rate or it grows but at a gradually decreasing rate). Finally, β_0 is the constant term, α_i is an unobserved random effect, correlated to the regressors X_{it}, which captures all unobserved time-invariant factors that affect Y_{it}, and ε_{it} is an idiosyncratic error that changes across time and units.

For the second part, we consider the following baseline model specification:

$$Y_{it} = \beta_0 + \beta_1 X_{i(t-1)} + \beta_2 X^2_{i(t-1)} + \beta_3 X_{i(t-2)} + \beta_4 X^2_{i(t-2)} + Z_i + u_{it} \qquad [2]$$

where Z_i is a set of relevant time-invariant variables at provincial level. As usual in the random-effect panel regression analyses, we assume that the α_i term is uncorrelated with the regressors X_{it} and it is included in the error term, so that $u_{it} = \varepsilon_{it} + \alpha_i$.

As a sensitivity analysis on the effect of Covid-19 pandemic on our dependent variable, model specifications illustrated in equation [1] and equation [2] are replicated replacing Covid-19 cases with Covid-related deaths. Estimations results of these alternative model specifications are provided in Appendix.

4. Results

The econometric results indicate that all the periods analyzed, except period Aug 2021 - Dec 2021, report a significant and positive relationship between RDC recipients and Covid-19 cases (Table 1). Looking at the magnitude of coefficients, the same extent of Covid-19 cases appears leading to a much higher number of RDC recipients during the first period. This evidence confirms that massive lockdown measures implemented by the Italian government to contrast the first wave of contagious strongly reduced the transaction costs related to the RDC claiming. Interestingly, even the Covid-19 cases reported at time t-2 seem to be significant in (positively) explaining the growth of RDC recipient households. The intertemporal effect of Covid-19 cases on the dependent variable is likely related to the fact that the RDC claim needs time (at least a couple of weeks) to become benefit receipt. To be noted, under this perspective, we should also consider that households may ponder for some time whether to claim for the RDC benefit once affected by a negative economic shock. Coefficients of the quadratic form of our both variables of interest, when statistically significant, present a negative sign during 2020, highlighting that the effect of Covid-19 cases consists of increasing RDC recipients but with decreasing marginal rates. In other words, the pandemic has led to a rise of RDC claims (and recipients), but this effect tends to taper off as Covid-19 cases gradually increase.

Table 1. Effects of growth of Covid-19 cases on RDC recipients (fixed-effects panel model)

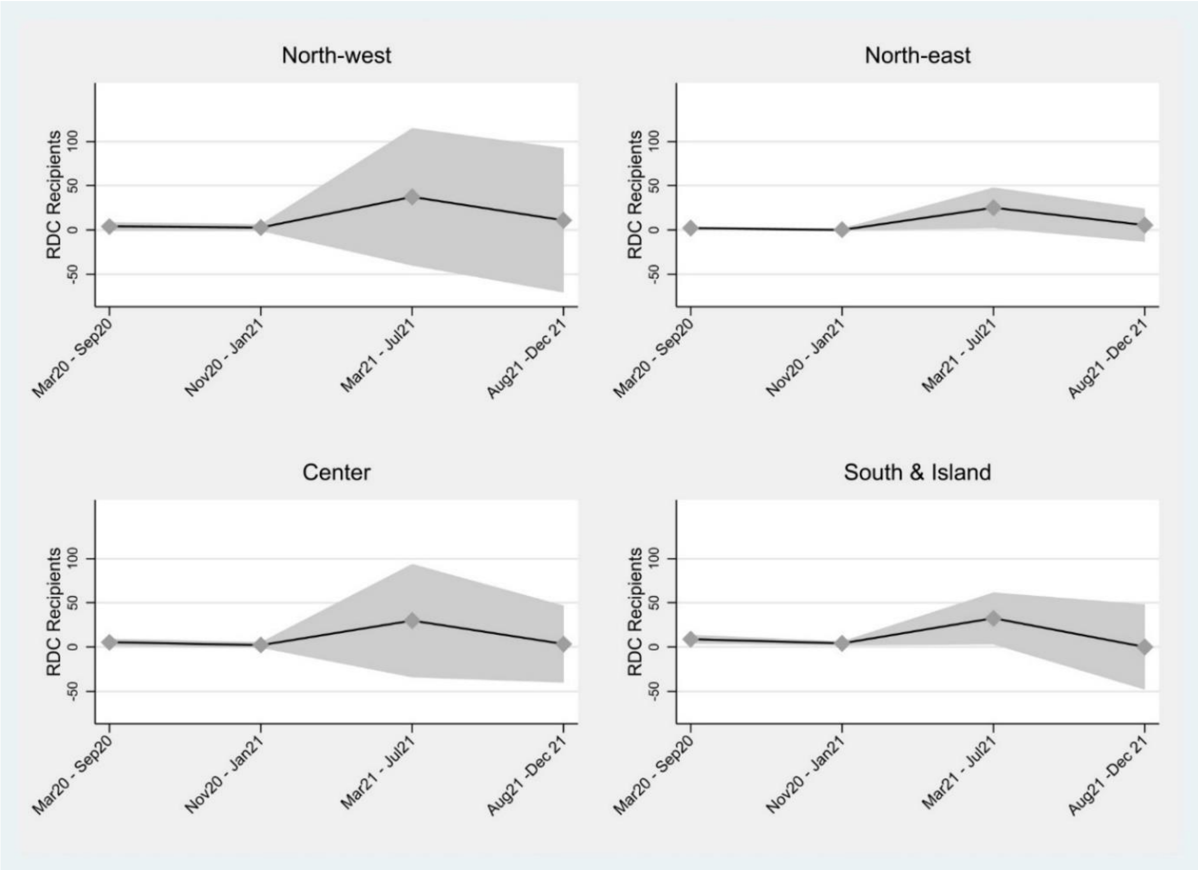
Variables	Mar 2020 - Sep 2020	Nov 2020 - Jan 2021	Mar 2021 - Jul 2021	Aug 2021 - Dec 2021
Cases (t-1)	1.021***	0.179***	0.922***	-0.212***
	(0.269)	(0.031)	(0.130)	(0.037)
$Cases^2$ (t-1)	-0.121***	-0.003***	-0.006***	0.001***
	(0.030)	(0.001)	(0.001)	(0.000)
Cases (t-2)	0.730***	0.070**	0.029	0.396***
	(0.124)	(0.030)	(0.079)	(0.059)
$Cases^2$ (t-2)	0.014	-0.001	0.001*	-0.002***
	(0.011)	(0.001)	(0.001)	(0.000)
Constant	42.229***	41.780***	11.200***	39.497***
	(0.355)	(0.249)	(2.148)	(1.473)
Average number of cases at time t-1	0.63	9.52	5.66	2.43
Average number of cases at time t-2	0.59	7.10	7.01	1.76
Observations	642	321	535	535
R-squared	0.196	0.514	0.808	0.328

*Notes: Standard errors clustered by Italian NUTS-3 level in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.1.*

The same considerations can also be extended for the third period (i.e. March-July 2021) except for the fact that, in this period, the effects at time t-2 lose significance. That could be explained by the adaptation of the RDC's administrative application systems to the situation of remote working, which overall engendered a speeding up of the application process. The period between August and December 2021 in contrast deviates from the expected results shown in the other three periods. It is likely that economic openings, mass vaccinations, and the new phase of active living with the virus has led to a reduction in the relationship between the pandemic and RDC applications, with the latter being more influenced by other factors.

Narrowing the analysis to the macro-regional level, the estimated confidence intervals confirm the significant and positive relationship between RDC recipients and Covid-19 cases in the first two periods (Figure 5). The third period is significantly greater than zero only for the North-East and the South of Italy, while the pandemic effect is always insignificant in the fourth period.

Figure 5. Confidence intervals of the overall Covid-19 cases effect by macro-region of residence



Notes: The figure shows the sum of the estimated coefficients multiplied by the average number of cases in each period by macro-region. The dark grey areas represent 90% confidence intervals.

When we replace the Covid-19 cases variable with the one reporting Covid-related deaths, our main considerations remain overall the same except for two points worth of mentioning (Table A3). First, the magnitude and significance of the coefficient over the period March 2020-September 2020 are stronger at time (t-2) than at time (t-1), differently from what recorded for the variable on Covid-19 cases. The second aspect concerns the magnitude of the effect of deaths at time (t-1) registered in the third period. The second aspect concerns the magnitude of the effect of deaths at time (t-1) recorded in the third period. Such a high coefficient is however counterbalanced by an equally high (and negative) coefficient of the quadratic form.

4.1. *A deepening on the role of territorial characteristics*

The random-effects model (Table 2) shows similar results to the previously described fixed-effects model. In particular, the significant and positive correlation between covid cases and RDC recipients is confirmed.

Table 2. Effects of growth of Covid-19 cases on RDC recipients (random-effects panel model)

Variables	Mar 2020 - Sep 2020	Nov 2020 - Jan 2021	Mar 2021 - Jul 2021	Aug 2021 - Dec 2021
Cases (t-1)	-0.158	0.177***	0.696***	-0.216***
	(0.157)	(0.031)	(0.118)	(0.039)
$Cases^2$ (t-1)	-0.017*	-0.003***	-0.004***	0.001***
	(0.010)	(0.001)	(0.001)	(0.000)
Cases (t-2)	1.008***	0.065**	0.103	0.410***
	(0.152)	(0.030)	(0.071)	(0.062)
$Cases^2$ (t-2)	-0.030***	-0.001	0.000	-0.002***
	(0.011)	(0.001)	(0.001)	(0.000)
Recipients in January 2020	1.217***	1.213***	1.348***	1.361***
	(0.028)	(0.042)	(0.055)	(0.048)
Share of foreign inhabitants	0.081	0.162	0.157	0.128
	(0.084)	(0.129)	(0.205)	(0.181)
Dependency ratio	-0.009	-0.053	0.210	-0.094
	(0.105)	(0.154)	(0.251)	(0.188)
Poverty index	0.021	0.029	0.322**	-0.025
	(0.059)	(0.101)	(0.138)	(0.114)
Unemployment rate	-0.057	-0.005	0.152	0.053
	(0.101)	(0.159)	(0.233)	(0.193)
Share of population living in a peripheral municipality	0.008	-0.006	0.067**	0.014
	(0.012)	(0.019)	(0.028)	(0.023)
Share of people with upper secondary education level	-0.006	-0.033	-0.004	-0.118
	(0.052)	(0.090)	(0.125)	(0.102)
Crimes	-0.051	-0.014	0.047	0.002
	(0.042)	(0.068)	(0.110)	(0.073)
Total mortality rate (per 10.000 inhabitants)	-0.000	-0.000	-0.000	-0.000**
	(0.000)	(0.000)	(0.000)	(0.000)
Women	0.035	0.041	-0.028	0.050
	(0.025)	(0.035)	(0.055)	(0.045)
Average household members	0.000***	0.000***	0.000	0.000***
	(0.000)	(0.000)	(0.000)	(0.000)
Constant	6.541***	6.970*	1.929	5.887
	(2.380)	(3.906)	(6.024)	(4.423)
Average number of cases at time t-1	0.63	9.52	5.66	2.43
Average number of cases at time t-2	0.59	7.10	7.01	1.76
Observations	642	321	535	535

*Notes: Standard errors clustered by Italian NUTS-3 level in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.1.*

Looking at coefficients of the other covariates, the variable with the larger effect and higher statistical significance is the one on RDC recipients in January 2020. In other words, territories that showed a greater use of the measure before the pandemic outbreak recorded a greater increase in applications. The importance of this proxy of the ‘territorial state dependence’ to the RDC benefit on our dependent variable appears however in line and supported by the literature on the topic (see for example Bhuller et al., 2017). The state dependence variable, which is correlated with the other demographic and socio-economic variables included in the model, likely leads to an underestimation of the latter, which in fact are not statistically significant in most cases and for all the periods examined. The only other covariate being significant (except for the third period) is the average household size. This evidence can be related, on the one hand, to a greater generosity of the RDC benefit for larger households and, on the other one hand, to the greater economic vulnerability featuring households with children in Italy (Del Boca and Mancini, 2013).

The same considerations are obtained by replacing the explanatory variable on the number of infections with that on the number of Covid-linked deaths (Table A4). Again, the positive correlation between deaths and RDC recipients is confirmed, demonstrating that, beyond the territorial use of the measure, the pandemic played a role in RDC claims.

4.2. Robustness checks

To test the robustness of our results we present below two particularly relevant robustness checks. The first robustness check assesses to what extent the spread of Covid-19 contagions have reduced transaction costs related to the RDC claiming mainly for a matter of economic loss or some further non-take-up dimension takes place in this case. As for the second check, considering that RDC recipients represent a subsample of those who actually claim for the RDC benefit, it assesses whether the trend of RDC recipients effectively reflects the greater propensity to claim for social benefit during pandemic.

There may be a legitimate suspicion that the trend of RDC recipients is strongly (or even exclusively) explained by the worsening economic conditions caused by the pandemic. This would undermine the starting hypothesis that identifies transaction costs as one of the drivers of RDC recipients' trend. We therefore contrasted the growth rates of RDC recipients (between January 2020 and December 2021) with the growth rates of taxable incomes between 2019 and 2020 (declared to fiscal authorities in 2020 and 2021 respectively). The comparisons are again performed on a provincial basis. Figure A6 shows the economic decline in the country, as most provinces exhibit a negative growth in taxable income. However, focusing on the correlation between the taxable income trend and RDC recipients one, we notice that provinces with the highest increase of the RDC incidence on provincial population are not the ones with the worst performance in income trajectory. In fact, the trend line is almost flat and the slope seems rather to indicate an inverse relationship, namely that the provinces with the most declining taxable incomes are those where the number of recipients has increased the least. Therefore, this evidence suggests that the role of pandemic on the spread and the increase of RDC recipients goes beyond the Covid-related income loss, shedding light on the relevance of a reduction of non-monetary transaction costs (e.g. stigma, conditionality, administrative barriers) in this case.

As explained before, the second robustness check tests our methodological decision of adopting the number of RDC recipients at territorial level as dependent variable rather than the number of RDC claimants. Table A5 presents the results of this check replicating Table 1 with the alternative dependent variable. Clearly, the positive and significant relationship between Covid-19 cases and the RDC spread is confirmed also in this case. Moreover, the magnitude of coefficients is higher than the one reported in Table 1, and the relationship of interest holds for the period August-December 2021 as well. To be noted, the same patterns emerge when we use Covid-related deaths instead of cases and when we estimate the random-effects panel model (more details are available upon request). Despite our results are confirmed when extending the analysis to the whole number of RDC claimants, we preferred focusing on the number of RDC recipients for the main analysis for three different reasons. First, looking at claimants, we may have a number of duplications as some households may have applied multiple times due to errors in documentations or hoping of being eligible for the benefit in a different moment of time. Second, the 18-month limit of the RDC receipt falling in October 2020 (see Section 3) led many existing recipients to reapply. Finally, claimants may also be non-eligible to the RDC benefit, so that we would include households with different economic conditions.

5. Conclusions and policy implications

We study the impact of the pandemic on population behaviors regarding the social assistance claiming, focusing on the case of the Italian minimum income scheme measure. Results show a significant and positive relationship between RDC recipients and Covid-19 cases trends, suggesting that the number of RDC recipients increased during periods of lockdown, which have been particularly pronounced during the first wave of contagions. This evidence confirms that the massive mobility restrictions implemented by the Italian government to counteract the virus spread strongly affected the economic well-being of households, and then reduced – as a collateral effect – the transaction costs associated with applying for the RDC benefit. Our results also appear robust to a change of Covid-19 spread proxy at the territorial level (Covid-19 cases vs Covid-related deaths) and to the consideration of relevant covariates directly influencing the RDC claim at the territorial level, so that the main conclusions of our study overall hold.

Interestingly, the pandemic impact on the minimum income scheme claim extends beyond the worsening of economic conditions caused by the same pandemic. Analyzing the correlation between taxable income trends and RDC recipients, it is evident that NUTS-3 regions with the highest increase in the incidence of RDC recipients are not necessarily those reporting the worst income trajectories during the first year of the pandemic. This sheds light on the relevance of reducing non-monetary transaction costs, such as social stigma, red tape bureaucracy, and conditionality. The advent of the pandemic temporarily interrupted conditionality measures related to the RDC benefit, forced both claimants and public administration offices to use digital/online tools and faster communication channels (e.g. telephone or e-mails), and overall reduced the afraid/fear of asking for help. As a final result, reducing transaction costs to claim for an important social transfer like the RDC one likely had a positive impact on the well-being of households in financial distress, especially in some regions of the country.

In conclusion, social support policies should be designed to have simpler application procedures that avoid excessive bureaucratic costs for potential recipients. More complex yet equally important would be addressing social perception and stigma. Awareness campaigns, positive communication that avoids stigmatizing prejudices, conscious training of social service operators, or other similar strategies may help decline the stigma associated with social assistance, encouraging those in need to seek support without fear of judgment. Finally, it is essential to ensure some level of flexibility in eligibility criteria and conditionality measures generally related to minimum income schemes during periods of crisis. Even better, this flexible decision should fall to regional authorities which better know – at least in theory – the characteristics and needs of their populations and labor markets. Looking at the Italian case, for instance, the temporary suspension of the mandatory active job search seems to have led a share of the eligible but non-recipient population to claim the cash benefit. While this outcome is already a positive one in 'normal' times, because it decreases the non-take-up issue and ensures that more people have a better standard of living, it turns out to be of great importance during a dramatic economic crisis. Welfare policies should therefore be designed to adapt to emerging needs, ensuring as possible that support is accessible to those who – even temporarily – need it.

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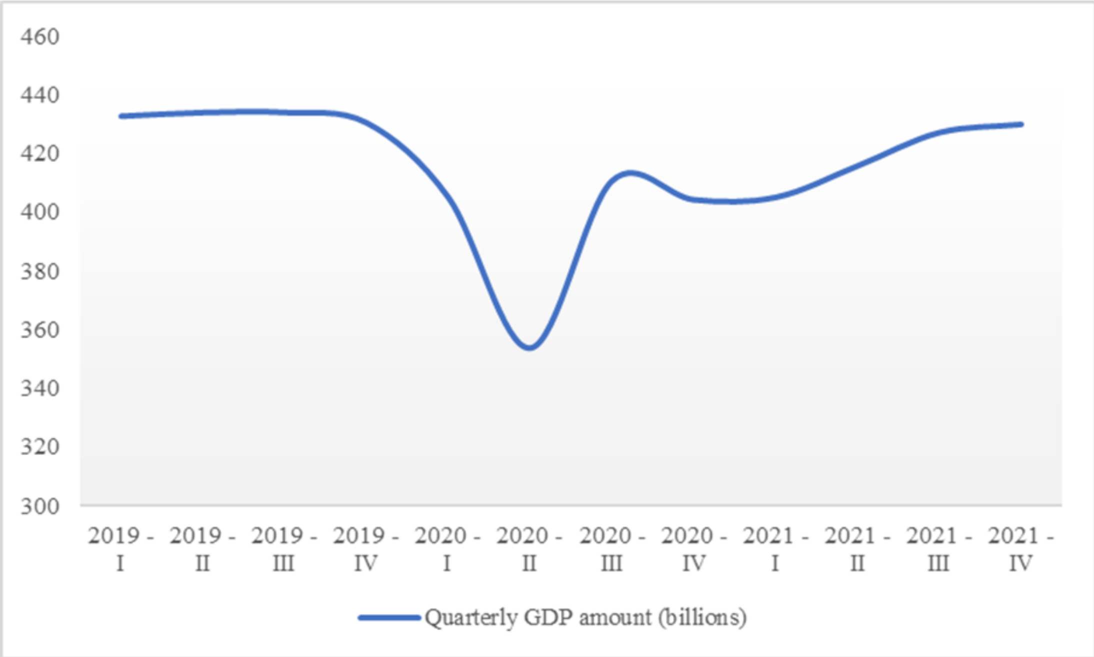
Appendix

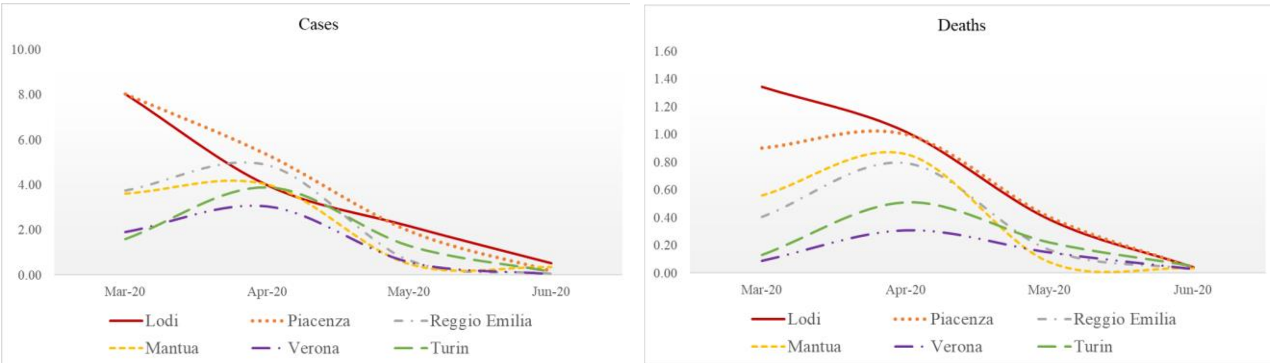
Figure A1. Quarterly GDP amount time trend between 2018 and 2021



Source: Elaboration of the authors on Istat data (2021).

Figure A2. Trend of Covid-19 cases and deaths (per 1,000 inhabitants) during the first stage of pandemic in a selection of Italian provinces

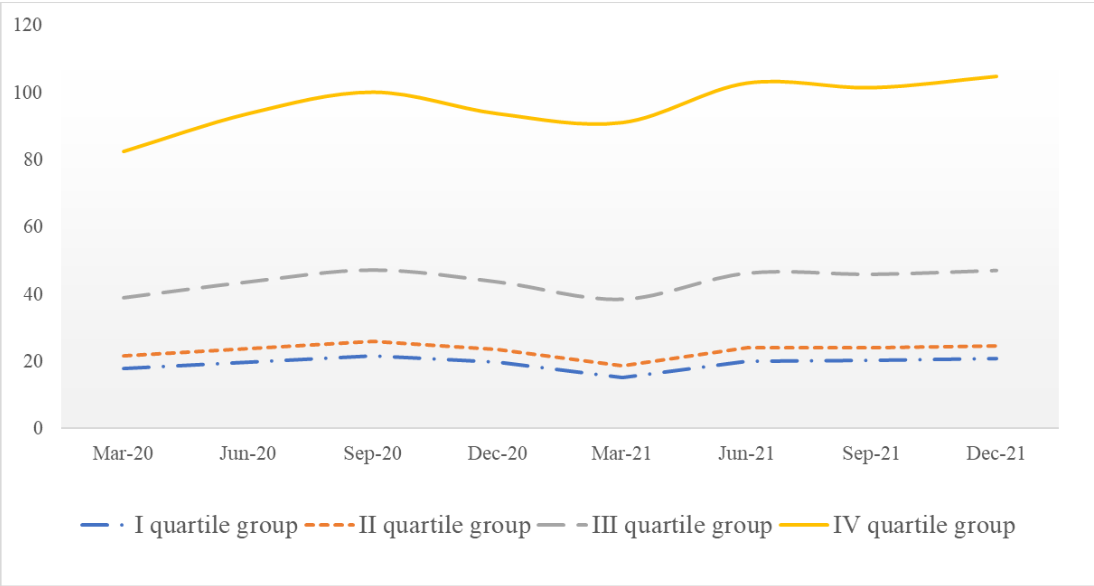
Panel A *Panel B*



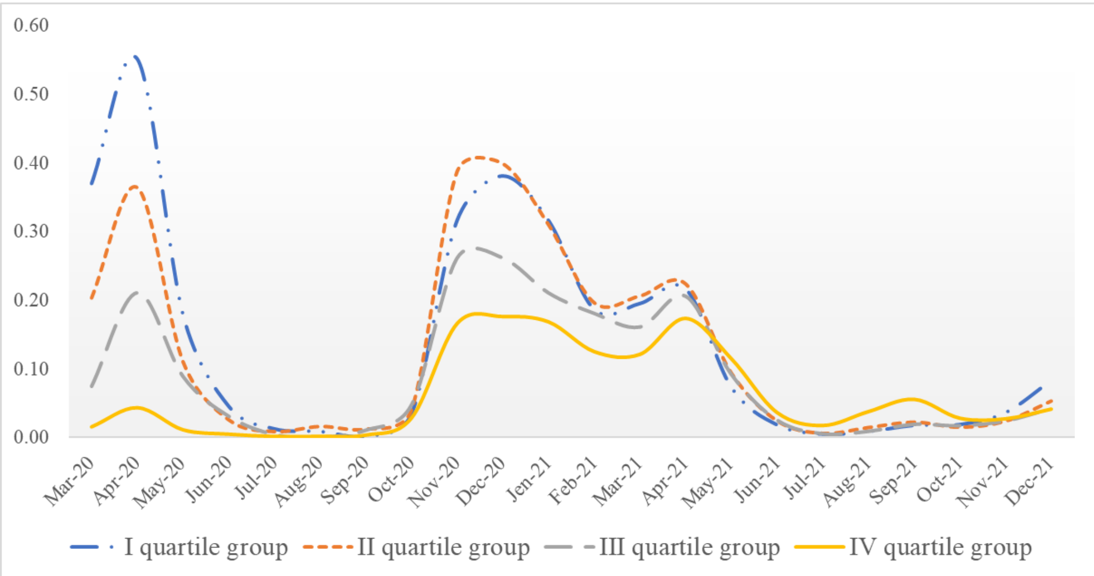
Source: Elaboration of the authors on Civil Protection Department data (2021). Notes: The number of cases and deaths reported is the one collected on the 28th day of the month. Notes: The provinces of Lodi, Mantua and Turin belong to the North-west of Italy, while Piacenza, Reggio Emilia and Verona to the North-east.

Figure A3. RDC recipients per 1,000 households (Panel A) and deaths due to Covid-19 per 1,000 inhabitants (Panel B) in Italian provinces divided by unemployment rate quartile group

Panel A



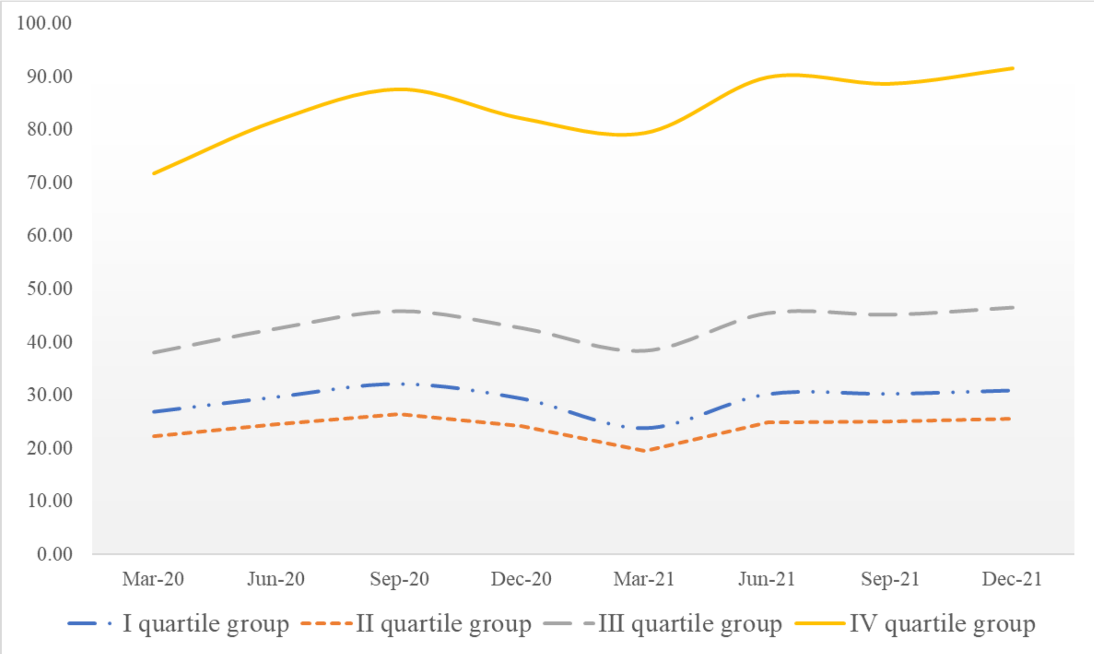
Panel B



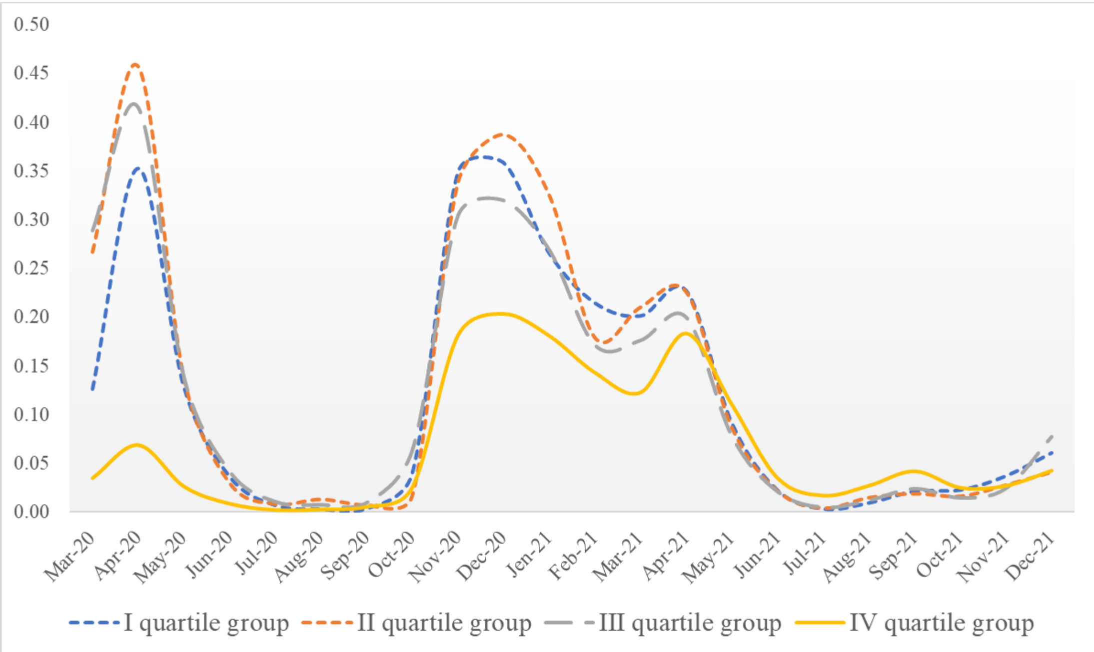
Source: Elaboration of the authors on INPS data (2021), Civil Protection Department data (2021), and ISTAT data (2019).

Figure A4. RDC recipients per 1,000 households (Panel A) and deaths due to Covid-19 per 1,000 inhabitants (Panel B) in Italian provinces divided by dependency ratio quartile group

Panel A

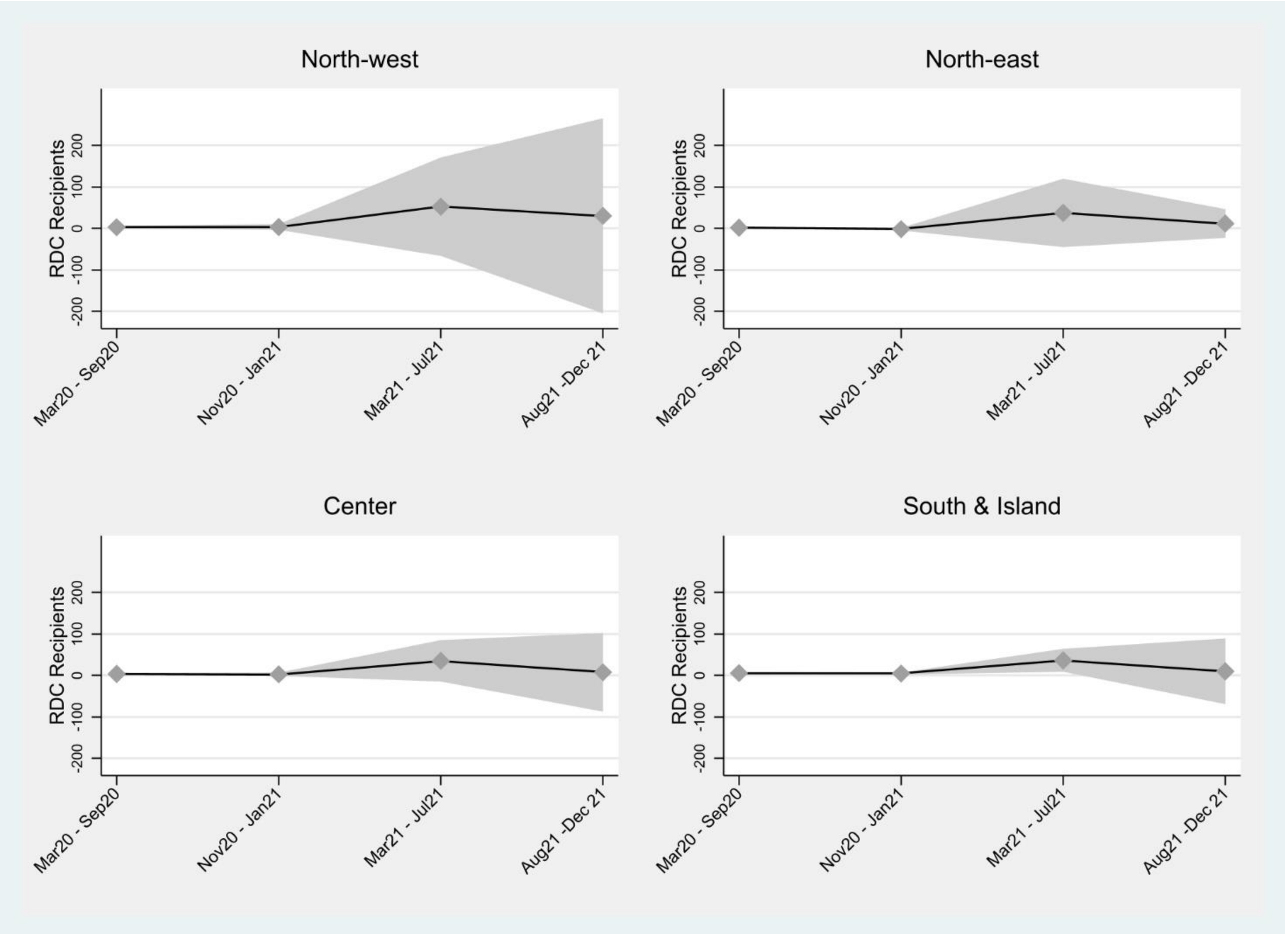


Panel B



Source: Elaboration of the authors on INPS data (2021), Civil Protection Department data (2021), and ISTAT data (2019).

Figure A5. Estimates confidence intervals on deaths by macro-region.



Notes: The figure shows the sum of the estimated coefficients multiplied by the average number of Covid-related deaths in each period by macro-region. The dark grey areas represent 90% confidence intervals.

Figure A6. Relationship between growth rates of RDC recipients and taxable incomes by Italian province

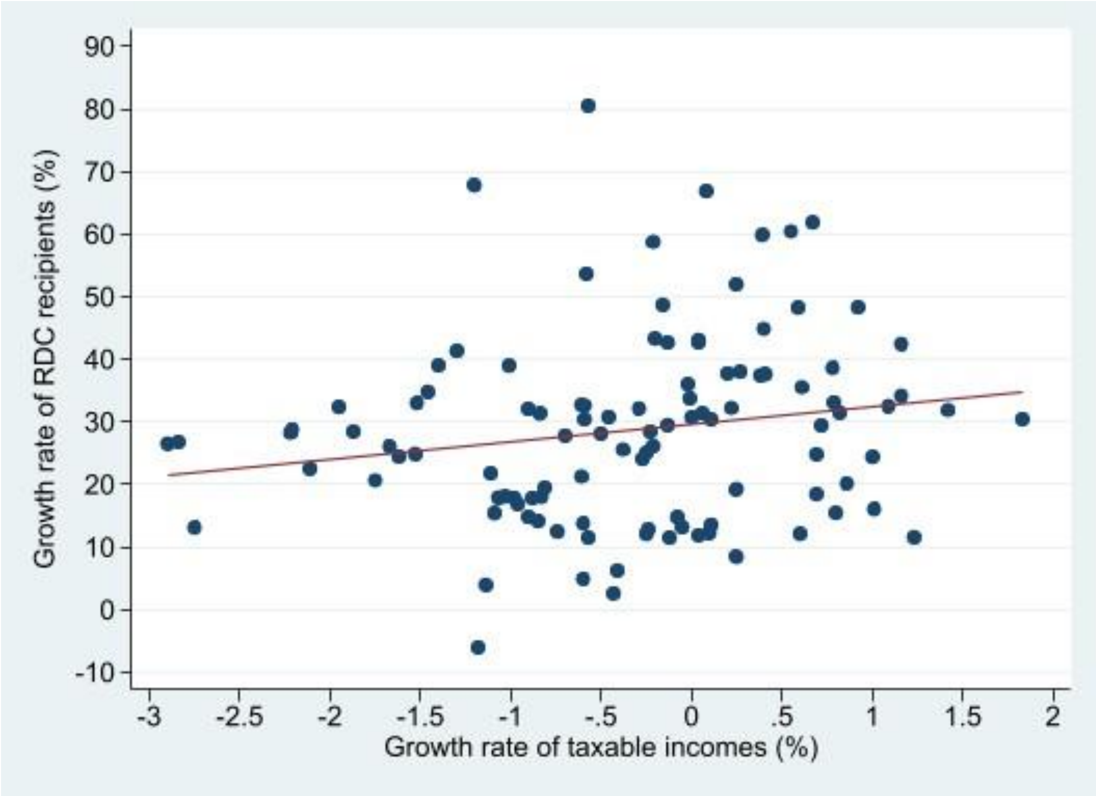


Table A1. List of variables used, including definition, source, and reference year

Variables	*Source (Year of reference)*	*Definition*
RDC recipients (dependent variable)	Osservatorio sul Reddito e Pensione di Cittadinanza: Italian National Social Security Institute (INPS) (2021)	Monthly trend in the number of RDC recipient households per 1,000 households at the provincial level
RDC applicants (dependent variable)	Osservatorio sul Reddito e Pensione di Cittadinanza: Italian National Social Security Institute (INPS) (2021)	Monthly trend in the number of RDC applicant households per 1,000 households at the provincial level
Cases	Civil Protection Department (2021)	Monthly trend in the number of people with Covid-19 infection per 1,000 inhabitants at the provincial level
Deaths	Civil Protection Department (2021)	Monthly trend in the number of people deceased with Covid-19 infection per 1,000 inhabitants at the provincial level. As this information is available at the regional level only, the variable is calculated for each province weighting regional Covid-19 deaths by its share of regional Covid-19 cases
Women	National institute of statistics ISTAT (2019)	Share of women at the provincial level
Average household members	National institute of statistics ISTAT (2019)	Average household members at the provincial level
RDC Recipients Jan-2020	Osservatorio sul Reddito e Pensione di Cittadinanza: Italian National Social Security Institute (INPS) (2021)	RDC recipient households per 1,000 households at January 2020 at the provincial level
Foreign inhabitants	National institute of statistics ISTAT (2019)	Share of foreign inhabitants on total population at the provincial level
Dependency ratio	National institute of statistics ISTAT (2019)	Age-population ratio of those not in the labor force to those in the labor force (i.e. aged 18-65) at the provincial level
Poverty rate	Ministry of Economics and Finance (2018)	Share of taxpayers declaring a taxable income lower than 10,000 € on total taxpayers at the provincial level
Unemployment rate	National institute of statistics ISTAT (2019)	Unemployment rate (people aged 15-74) at the provincial level
PM_pop	Ministry of Economic Development (2014)	Share of population living in a peripheral municipality at the provincial level
High school graduation rate	National institute of statistics ISTAT (2011)	Share of people aged 19 or more attained the upper secondary education level at the provincial level
Crimes	Ministry of Interior (2018)	Number of crimes at the provincial level
Mortality rate	National institute of statistics ISTAT (2017)	Total mortality rate per 10.000 inhabitants at the provincial level
Taxable incomes	Ministry of Economy and Finance (2020 and 2021)	Aggregate taxable income at provincial level

Table A2. Monthly percentage increases in RDC recipients per macro areas

	North-West	North-East	Center	South
Sep-19	-1,7%	-2,9%	-2,0%	-1,5%
Oct-19	3,4%	3,5%	3,2%	3,9%
Nov-19	3,4%	2,8%	3,5%	4,0%
Dec-19	3,0%	2,9%	2,6%	2,6%
Jan-20	2,1%	1,9%	1,6%	1,4%
Feb-20	-0,5%	-1,9%	0,1%	1,6%
Mar-20	-0,5%	-1,9%	0,1%	1,5%
Apr-20	3,6%	3,2%	3,9%	5,0%
May-20	3,5%	3,0%	4,5%	4,8%
Jun-20	4,3%	3,4%	4,9%	4,2%
Jul-20	4,6%	3,3%	4,5%	3,3%
Aug-20	4,1%	3,2%	3,6%	2,5%
Sep-20	2,7%	2,1%	2,4%	1,4%
Oct-20	-23,1%	-22,4%	-24,3%	-29,1%
Nov-20	16,8%	13,3%	20,9%	29,3%
Dec-20	5,0%	3,7%	4,7%	2,5%
Jan-21	-0,4%	-2,1%	0,3%	2,5%
Feb-21	-31,8%	-35,5%	-26,2%	-14,1%
Mar-21	15,7%	18,5%	15,1%	10,3%
Apr-21	11,9%	16,2%	10,4%	5,4%
May-21	9,8%	12,3%	9,6%	4,6%
Jun-21	4,0%	3,6%	4,3%	2,7%
Jul-21	3,3%	3,2%	3,1%	2,1%
Aug-21	-0,6%	-0,1%	-0,9%	-1,6%
Sep-21	-1,9%	-1,8%	-1,7%	-1,6%
Oct-21	0,7%	0,7%	1,8%	1,7%
Nov-21	0,7%	1,0%	0,2%	0,4%
Dec-21	0,4%	0,9%	0,8%	1,1%

Source: Elaboration of the authors on INPS data (2021).

Table A3. Effects of growth of Covid-related deaths on RDC recipients (fixed-effects panel model)

Variables	Mar 2020 - Sep 2020	Nov 2020 - Jan 2021	Mar 2021 - Jul 2021	Aug 2021 - Dec 2021
Deaths (t-1)	2.011**	5.401***	28.687***	-2.312
	(0.775)	(0.750)	(3.486)	(3.091)
$Deaths^2$ (t-1)	-2.336***	-1.423***	-4.201***	0.443
	(0.683)	(0.235)	(0.674)	(0.437)
Deaths (t-2)	6.006***	-0.668	0.338	15.210***
	(0.689)	(0.756)	(2.474)	(2.675)
$Deaths^2$ (t-2)	-0.585***	0.264	0.470	-2.132***
	(0.200)	(0.209)	(0.459)	(0.500)
Constant	43.495***	41.638***	7.540***	29.887***
	(0.137)	(0.420)	(2.360)	(2.521)
Average number of deaths at time t-1	0.10	0. 21	0.13	0.02
Average number of deaths at time t-2	0.12	0.11	0.17	0.02
Observations	642	321	535	535
R-squared	0.156	0.331	0.764	0.307

*Notes: Standard errors clustered by Italian NUTS-3 level in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.1.*

Table A4. Effects of growth of Covid-related deaths on RDC recipients (random-effects panel model)

Variables	Mar 2020 - Sep 2020	Nov 2020 - Jan 2021	Mar 2021 - Jul 2021	Aug 2021 - Dec 2021
Deaths (t-1)	-1.888***	5.550***	15.617***	-3.186
	(0.688)	(0.782)	(3.856)	(3.051)
Deaths^2 (t-1)	-0.358	-1.418***	-3.121***	0.703*
	(0.257)	(0.237)	(0.762)	(0.422)
Deaths (t-2)	6.307***	-1.258*	8.717***	16.085***
	(0.826)	(0.697)	(2.769)	(2.609)
Deaths^2 (t-2)	-1.092***	0.379*	-0.427	-2.802***
	(0.266)	(0.197)	(0.538)	(0.423)
Recipients January 2020	1.216***	1.224***	1.408***	1.392***
	(0.028)	(0.043)	(0.059)	(0.049)
Share of foreign inhabitants	0.077	0.149	0.098	0.104
	(0.081)	(0.124)	(0.229)	(0.175)
Dependency ratio	-0.006	-0.040	0.098	-0.114
	(0.103)	(0.159)	(0.310)	(0.192)
Poverty index	0.008	0.067	0.519***	0.093
	(0.058)	(0.099)	(0.169)	(0.107)
Unemployment rate	-0.058	-0.024	0.064	0.044
	(0.100)	(0.158)	(0.244)	(0.190)
Share of population living in a peripheral municipality	0.008	-0.004	0.066*	0.031
	(0.012)	(0.019)	(0.036)	(0.024)
Share of people with upper secondary education level	-0.013	0.005	0.119	-0.074
	(0.051)	(0.095)	(0.182)	(0.113)
Crimes	-0.045	0.022	0.274**	0.080
	(0.041)	(0.067)	(0.132)	(0.076)
Total mortality rate (per 10.000 inhabitants)	-0.000	-0.000	-0.000	-0.000***
	(0.000)	(0.000)	(0.000)	(0.000)
Women	0.033	0.039	-0.028	0.014
	(0.025)	(0.034)	(0.062)	(0.043)
Average household members	0.000***	0.000***	0.000**	0.000***
	(0.000)	(0.000)	(0.000)	(0.000)
Constant	6.400***	7.014*	4.300	3.041
	(2.361)	(3.895)	(7.624)	(4.557)
Average number of deaths at time t-1	0.10	0. 21	0.13	0.02
Average number of deaths at time t-2	0.12	0.11	0.17	0.02
Observations	642	321	535	535

*Notes: Standard errors clustered by Italian NUTS-3 level in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.1.*

Table A5. Effects of growth of Covid-19 cases on RDC claimants (fixed-effects panel model)

Variables	Mar 2020 - Sep 2020	Nov 2020 - Jan 2021	Mar 2021 - Jul 2021	Aug 2021 - Dec 2021
Cases (t-1)	1.649***	0.013	0.947***	1.173***
	(0.360)	(0.061)	(0.157)	(0.233)
$Cases^2$ (t-1)	-0.175***	-0.003**	-0.006***	-0.005***
	(0.040)	(0.001)	(0.001)	(0.001)
Cases (t-2)	0.970***	0.913***	0.064	0.676***
	(0.158)	(0.108)	(0.088)	(0.226)
$Cases^2$ (t-2)	0.029*	-0.013***	0.001	-0.002
	(0.016)	(0.002)	(0.001)	(0.002)
Constant	70.241***	84.393***	76.670***	30.529***
	(0.481)	(0.621)	(2.795)	(6.062)
Average number of cases at time t-1	0.63	9.52	5.66	2.43
Average number of cases at time t-2	0.59	7.10	7.01	1.76
Observations	642	321	535	535
R-squared	0.296	0.568	0.774	0.725

*Notes: Standard errors clustered by Italian NUTS-3 level in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.1.*

Chapter 2:

Surnames in local newspapers and social mobility *

Abstract

This study aims to investigate social mobility by introducing an innovative data source: newspapers. The core hypothesis posits that frequent mentions in newspapers are indicative of greater social relevance. Through the collection of local newspapers data in the municipality of Modena, Italy, from 1921 to 2011, we examine variations in the relative representation of surnames in newspapers respect to their presence in registers data. Results suggest that surnames in newspapers are not a random sample of the population, supporting the assumption that they reflect social significance. Surnames belonging to privileged groups exhibit a higher representation compared to other social groups. This higher relative representation seems to be transmitted to future generations and converges over time toward the mean, with variations depending on the considered high-status group. This kind of analysis could contribute to identifying different mobility patterns at the local level and represent a useful alternative when established data sources, especially income data, are not available.

Keywords: newspapers, social mobility, surnames, elites

JEL classification codes: J62, N34

* This chapter is co-authored with Massimo Baldini (University of Modena and Reggio Emilia).

1. Introduction

The study of social mobility between generations has always stimulated academic debate (see among others Ganzeboom et al., 1991). However, the difficulty in finding data linking generations over time has held back for a long time the development of research on this subject. It is only in recent decades that the empirical literature on intergenerational mobility has made considerable progress, mainly due to the availability of new and better data[1] linking adjacent generations (Solon, 2018). Social mobility over long periods has received less attention from scholars because of data limitations (Solon, 1999; Black and Devereux, 2010). This has pushed in parallel to the search for alternative methods and data to overcome this obstacle. In particular there is a growing interest in the exploitation of surnames in studying social mobility, used as pseudo-links connecting households of different generations (Barone and Mocetti, 2021). The present work builds on this literature by introducing a novel data source for the analysis of social mobility: newspapers.

We investigate the relative representation (see Clark and Cummings, 2015) of surnames from different social groups in newspaper articles compared to their presence in the population, assuming that individuals frequently mentioned in newspapers hold social relevance. If privileged groups (nobility, cultural elites, etc.) are more frequently represented in newspapers than in population registers, this confirms the hypothesis that newspaper surnames are not randomly selected but carry traits of social relevance.[2] Therefore, if this over-representation persists over time, it indicates the transmission of power across generations.

Several features derived from these data can be used to analyse social mobility over time, bearing in mind that the informational content of a surname can vary greatly depending on the context and type of the article in which it appears. For example, many of the surnames that appear in the political section of a newspaper correspond to people who have power in collective decision-making processes; in crime-related articles, the surnames of both the victims and the criminals (often of lower socio-economic status) appear; those who advertise their products/services may be considered richer

[1] These types of data mainly refer to: (i) cross-sections with retrospective surveys on the social status of parents; (ii) panel surveys with detailed and repeated information on the social and economic status of household components followed over extended periods of time; (iii) fiscal/administrative data linking economic information between parents and children (Mooi-Reci, 2020).

[2] The hypothesis that the likelihood of being quoted in newspapers is associated with higher social status has already been examined in other studies. For instance, the literature studying obituaries emphasises that they reveal the influence of dominant Western elites, groups linked by common origins and educational patterns (Matsuda, 1996; Fowler and Bielsa, 2007). Ban et al. (2019) demonstrate that the volume of press coverage devoted to political actors or offices helps to indicate their actual power.

than the average member of the reference population, just like those who spend themselves on charity work or who can afford to buy space for an obituary.

To ascertain whether there have been variations in the relative representation (RR) of surnames over the years, the surnames mentioned in local newspapers in the municipality of Modena, Italy, from 1921 to 2011 are taken into account, together with registry data on the spread of surnames in the general population. By way of explanation, if there is an over-representation of certain surname groups in pages dealing with local politics compared to their registry distribution, then that surname group will most likely belong to individuals/families exercising political power at the local level. If this over-representation persists over time, then power is transmitted from one generation to the next.

As with other papers that exploit the information content of surnames, it is not possible to state with certainty that a surname observed today is part of the direct descendants of the same surname found in previous decades. But, in line with Barone and Mocetti (2021), by focusing on the local rather than the national level, more precise links between generations are produced. Moreover, the strong heterogeneity and ‘localism' of Italian surnames further reinforce the quality of these pseudo-links. Additionally, when adjusting for rare surnames in the population, the strength of these connections becomes even more apparent.

To the best of our knowledge, the contribution of this work to the existing literature is threefold. We are the first to propose the use of surnames appearing in newspapers to study social mobility and power transmission of elites. Moreover, while the literature on long-run social mobility has traditionally focused on income, education, and occupation outcomes (Solon, 2018), we implement an analysis that is rather concerned with the transmission of authority, power and social influence. Finally, the employment of newspapers as a source of historical data has mainly been developed in the United States and the United Kingdom, while in Europe, due to the lower presence of accessible digitised archives, it is a tool with wide scope for scientific research.

Our findings indicate that surnames in newspapers provide independent informative content and are not randomly extracted from the population. We find support for the hypothesis that surnames appearing in newspapers have particular social significance; those associated with higher social status groups are more represented compared to the rest of the population. Surnames linked to socially significant groups in the past tend to maintain a greater relative representation in newspapers over time, highlighting the transmission of social status across generations. This transmission regresses toward the mean at varying rates depending on the social group examined.

The rest of the paper is organised as follows. In Section 2, a brief review of the research on newspapers and on surnames (in particular in the study of social mobility) is proposed. Section 3 explains the data used. Section 4 reports the method of analysis and the observed results, while two robustness checks are proposed in section 5. Finally, Section 6 concludes.

2. Literature review

2.1. The use of surnames in economic research and the study of social mobility

A person's surname can tell a lot about the social status of the ancestors and the geographical origin of the family (e.g., the surnames Le Boulanger in France, Becker in the United States and Bäcker in Germany link their owners to their ancestors' occupation as bakers). Furthermore, the information conveyed by surnames may influence the way a person is judged, a fact that economic research has exploited to study discrimination in various fields, for example in the labour or in the rental markets. For instance, Pascual et al. (2015) find that popular surnames receive a higher and more positive evaluation than uncommon and infrequent ones in the labour market. Noble-sounding surnames more frequently hold managerial positions than other surnames (Silberzahn & Uhlmann, 2013). A person with an African American-sounding name is less likely to be called back for a job interview than a person with a white-sounding name (Bertrand & Mullainathan, 2004), and anonymous application procedures increase the chances of both women and individuals of non-Western origin of advancing to the interview stage (Åslund & Skans, 2012). Many studies have also found signs of discrimination in rental housing markets, where foreign-sounding names that have shown interest in the ads are less likely to be contacted by flat owners (Ahmed and Hammarstedt 2008).

As an alternative method of measuring intergenerational social mobility, the use of surnames has received increasing interest. Collado et al. (2012), having data on the socio-economic status of individuals in two Spanish regions at the end of the 19th and the end of the 20th century, use pseudo-links obtained from surnames to link ancestors to descendants. They find that having a high level of education and belonging to a high socioeconomic group is still influenced by the socioeconomic status of the great-great-grandparents. Guell et al. (2015) estimate a correlation of educational outcomes of 0.60 for Catalonia in Spain. Clark (2014) finds that in a wide set of examined countries multigenerational mobility follows an autoregressive process with a high and time-invariant persistence rate. Barone and Mocetti (2021) go even further back in time, exploiting a unique dataset linking the status of Florentines in 1427 (derived from a tax census) to that of their likely descendants (with the same surnames) in 2011. They find that the elasticity of earnings is positive and statistically significant between generations nearly 600 years apart. They also show evidence of even stronger transmission of real wealth and persistence in some elite professions. None of these studies, however, made use of surnames taken from newspapers.

2.2. Newspapers

Newspapers can be the source of an incredibly vast amount of easily accessible historic information (Hansen, 2004). In fact, their pages contain the historical memory collected day by day of major political, cultural, economic, scientific, and other events (Tosh, 2010).

Hanlon and Beach (2022) highlight that research in economics, before the arrival of large-scale digitized historical newspaper databases, has predominantly focused on the use of newspapers: (i) as a way of measuring a certain type of treatment in order to construct a key explanatory variable; (ii) as the basis of many price series and financial and commodity markets studies. They also emphasise that the bulk of existing economic studies have focused on data from the United States or the United Kingdom, where most of the digital archives are located. A very promising research perspectives is therefore to expand the use of newspapers outside this geographical area.

Newspaper data can be applied to a wide range of topics, but so far only some of them have been explored. Following the outbreak of the Covid-19 emergency, for example, a theme that has received special emphasis is the study of the 1918-19 influenza pandemic. Markel et al. (2007) demonstrate a strong association between the early and stratified implementation of non-pharmacological interventions such as school closures, isolation or quarantine and the mitigation of the negative consequences of the influenza pandemic in the United States. In addition to administrative data, the authors use information from newspapers to verify the type and date of intervention. Again, based on data from local US newspapers, Ager et al. (2022) investigate the impact of school closures to prevent the spread of influenza on long-term school outcomes.

The study of collective action with newspaper data is another topic that has become commonplace in recent decades. Multiple types of collective action, from racial violence protests to various other types of social movements, have been analysed through newspaper-based event data (for a review see Earl et al., 2004).

Through data obtained from newspapers, it is also possible to monitor economic activity in real time (Shapiro et al., 2020; Aguilar et al., 2021). Textual analysis and news sentiment are indeed exploited to measure in which direction the economy is moving. When journalists write more positive words, then the economy is experiencing an upward trend and vice versa. Somewhat related to news sentiment is the study by Gentzkow et al. (2006), who construct a historical index of corruption based on mentions of 'fraud' and 'corruption' reported in US newspapers.

3. Data

The analysis incorporates different data sources. The primary one is based on a dataset covering surnames mentioned in local newspapers in the municipality of Modena, Italy, spanning from 1921 to 2011. Since digitized archives of these newspapers are not available, data were collected manually by extracting surnames from physical articles and recording them in a database categorized by year, section (for further details, refer to later), and newspaper title. This entailed the need to restrict the survey to specific time intervals. Arbitrarily, surnames were gathered for one year per decade (1921, 1931, ..., 2011), considering the first three days of each month and the entire month of March.

It is important to collect data from different newspapers at the local level since media coverage is not independent of the biases of the newspapers themselves and the preferences of the readers (Gentzkow and Shapiro 2010; Larcinese et al 2011). There have been three local newspapers active in the territory of Modena in the reference period. Including all three newspapers reduces this bias.

- *La Gazzetta di Modena*[3], the main daily newspaper in the city of Modena, active throughout the time span analyzed. It is a moderate, generalist daily, not politically aligned (except of course during the fascist period).
- *L'Unità (local edition of Modena),* available for the years 1951-1961-1971-1981-1991 of our sample. Historical left-wing daily linked to the Italian Communist Party. Over the years it has gradually embraced more moderate and reformist positions following the evolution of its reference party.
- *Il Resto del Carlino (local edition of Modena)*, used for the years between 1961 and 2011. It is a generalist daily, historically affiliated with the agrarians of Emilia and the sugar industrialists during the first decades of the XX century and usually supporting centre-right parties after the second world war.

In order to distinguish the mentions according to the different spheres of influence of individuals and the type of article, surnames have been classified in pre-established sections: Advertising, Business, Charity, Public Events & Civil Society , Crime News & Court Reporting , Obituaries, Local Politics,

[3] The newspaper has gone through several name changes: La Gazzetta di Modena (1947-1953); Gazzetta dell'Emilia (1953-1967); Gazzetta di Modena (1968-1977); Nuova Gazzetta di Modena (1981-present).

Religion, Science & Technology. Surnames appearing in articles which refer to non-local (for example national politics) and sports events are excluded. Table 1 shows the number of mentions per year and per section. Additional information and statistics on this and the other datasets used in the paper can be found in the appendix tables.

The second dataset concerns the distribution of surnames within the population. This type of information is crucial as it allows us to compare how frequently a surname is mentioned in newspapers relative to its presence in the general population. We have the surnames of household heads in 1936, and of all residents in Modena in 1981 and 2001 from the Italian population registries 8th, 12th, and 14th.

The third dataset is essential for establishing a robust tool to assess newspapers' capacity in observing movements in social mobility. This dataset contains information regarding the surnames of students who attended 'Licei' in the municipality of Modena. In Italy, the educational system offers three main types of high school pathways: Licei, technical institutes, and vocational schools. While the latter two pathways directly prepare students for entry into the workforce, Licei are primarily designed to provide preparation for further university studies. The information pertaining to students who attended Licei is particularly relevant in the context of one of the most established indicators in the study of social mobility: education. This stems from the expectation that the surnames of individuals belonging to privileged social groups will exhibit a higher relative representation within the registries of Licei compared to their distribution in the general population. This trend is particularly anticipated for past years when access to university education was a privilege reserved for a few, and only more affluent families could afford to support their children's studies.

The subsequent phase involved the identification of surnames belonging to various social groups. Among the elite groups, we included surnames associated with the Modenese nobility, surnames of Modena university professors, and surnames from the Jewish community. [4] The other groups, used as control samples during the analyses, were randomly selected from the first surnames in alphabetical order, as well as from common surnames found in the 1936 registry with limited representation in the 1921 newspaper (below the 60th percentile of the median). Information on these surname groups is contained in Table A8.

[4] The city of Modena has hosted a community of Italian Jews since the Middle Ages. The history of this social group has gone through very heterogeneous phases in Italy, but at the beginning of the XX century many of its members occupied important social positions (for example, the major of Modena in 1921 had Jewish origins).

Table 1. Collected surnames by year and section.

Section	1921	1931	1941	1951	1961	1971	1981	1991	2001	2011	Total
Advertising	355	662	363	754	755	311	184	324	325	77	4,110
Business	205	373	181	284	197	403	277	63	394	177	2,554
Charity	359	998	552	434	361	212	73	0	0	0	2,989
Public Events & Civil Society	212	211	548	665	446	964	762	854	2,334	2,271	9,267
Crime News & Court Reporting	529	294	413	902	101	402	262	151	268	230	3,552
Obituaries	55	43	384	245	82	282	152	135	287	212	1,877
Local Politics	531	467	476	578	671	790	804	1,228	1,447	1,605	8,597
Religion	0	32	20	58	49	40	63	79	87	72	500
Science & Technology	74	366	111	148	404	313	258	224	395	347	2,640
Total	**2,320**	**3,446**	**3,048**	**4,068**	**3,066**	**3,717**	**2,835**	**3,058**	**5,537**	**4,991**	**36,086**

Notes: The Crime News & Court Reporting section is a combination of the subsections' thief, victim, accidents, and judges-lawyers-police.

4. Methods of analysis and results

The fundamental premise of this study is that the information collected from newspapers can be used as a valuable complement to existing measures of social mobility. To test this hypothesis, a series of steps need to be followed. First and foremost, we need to conduct analyses on newspaper data. In particular, it is important to:

i) Determine whether the frequency of surnames in newspapers provides independent informative content and is not a random extraction from the population.
ii) Verify whether the surnames that appear most frequently in local newspapers have significant social relevance.

The second step involves analyzing the mobility of surname groups in newspapers over time. This phase should be accompanied by a comparison of the results obtained with those achieved through a well-established measure in the study of social mobility, which in our case is the transmission of surnames through education.

In all these phases, the key statistic on which we rely is the relative representation. This is the ratio between the share of surnames or groups of surnames present in a context where it is presumed that privileged positions are transmitted (such as in newspapers or Licei) and the share of those same surnames in the population during the specified period. In other words, the relative representation of a specific group of surnames s is:

$$Relative\ Rapresentation\ of\ s\ = \frac{Share\ of\ s\ in\ privileged\ context}{Share\ of\ s\ in\ general\ population}$$

We have only access to the frequency distribution of surnames in the general population from registry data for the years 1936, 1981, and 2001. This necessitates using the same registry for multiple decades as the denominator for calculating relative representation. This challenge is mitigated by the fact that the population composition changes slowly over time. Specifically, for the years 1921, 1931, 1941, and 1951, we rely on the 1936 registry. For the years 1961, 1971, and 1981, we use the 1981 registry, while for the years 1991, 2001, and 2011, we refer to the 2001 registry. Therefore, the relative representation of noble surnames in 1921 newspapers is:

$$RR\ of\ noble\ surnames\ in\ newspaper_{1921} = \frac{Share\ of\ noble\ surnames\ in\ newspaper_{1921}}{Share\ of\ noble\ surnames\ in\ registry_{1936}}$$

4.1. Checking newspaper data

The first check on the newspaper data concerns the potential independent informational content. It is natural for the most common surnames in the general population to be widely represented in newspapers. However, what is of critical is whether the distribution of these surnames in newspapers represents a random sample of the general population.

To address this issue, we calculate the relative representation for each surname and assess whether these values exhibit a correlation with their respective relative frequencies in the general population (Table 2). If a strong correlation is observed, it indicates that the surnames in newspapers are merely a random reflection of the general population composition and do not provide autonomous information. For surnames reported in newspapers in 1921, the correlation coefficient between their RR and their relative frequency in the population is -0.23. In 1981, the same correlation is -0.13. For the 200 most common surnames in the 1936 registry, the correlation in 1921 is exactly 0 (Table A9 in appendix). For the 200 most frequent surnames in the 1981 registry data, the correlation is -0.04. In 1921, the 200 most common surnames in the 1936 registry exhibit an average RR of 0.57, which is close to the overall average (0.63). This trend persists in subsequent years. Therefore, we can conclude that the surnames in newspapers do not result from a mere random extraction from the underlying population.

Table 2. Correlation between the RR of surnames in newspapers and their frequency distribution in general population.

	RR 1921	RR 1931	RR 1941	RR 1951	RR 1961	RR 1971	RR 1981	RR 1991	RR 2001	RR 2011
Frequency distribution registry 1936	-0.23	-0.22	-0.19	-0.24						
Frequency distribution registry 1981					-0.16	-0.16	-0.13			
Frequency distribution registry 2001								-0.21	-0.08	-0.14

Notes: Frequency distributions in general population refer exclusively to surnames present in newspapers during the reference year. Therefore, the -0.23 value for 1921 is the outcome of the correlation between the RR of surnames present in 1921 newspapers and their frequency distribution in the 1936 registry.

The second control questions whether there is a basis for the assumption that surnames appearing in local newspapers possess particular social significance. Naturally, not all surnames in this source belong to the local elite, but for a portion of them, this could be a plausible scenario. Once this point is established, observing the evolution of surname distribution in newspapers over time can provide insight into changes in the structure of elite groups. If the distribution of surnames belonging to privileged groups shows few changes between two time periods, it may suggest that influential positions have not undergone significant alterations, and vice versa.

To test the hypothesis that surnames in newspapers contain information about elite groups, it is necessary to compare the relative representation of surnames that unquestionably belong to a socially relevant group. We expect surnames belonging to such groups to, on average, have a higher RR in the newspaper dataset compared to the rest of the population.

Table 3 confirms these expectations. Surnames associated with socially relevant groups are more prominently represented in newspapers than in the general population, and this trend remains consistent over the analyzed years. For the four examined elite groups, their relative representation is higher than that of the general population. In almost all cases, their relative representation is also above 1, a value indicating equitable distribution of surnames in newspapers and registers. For instance, noble surnames are almost four times more present in newspapers than in registy in 1921. The same holds for university professors teaching in 1921, with a relative representation of 3.42. This value further increases for professors in 1931, with a representation seven times higher in newspapers than in the general population.

It is noteworthy that when the number of surnames belonging to a group is limited, distorted estimates and increased standard errors may occur, indicating considerable data variability. Therefore, it is a positive sign that the analysis remains robust for the group of Licei students, where surname diversity and numerosity are higher.

An intriguing aspect also arises in the comparison of relative representation between typically Jewish surnames in Modena and others. Despite the numerical limitations of such surnames potentially influencing estimates for this group, in 1921, the relative representation of Jewish surnames is 4.31, almost eight times higher than non-Jewish surnames (0.59). In 1931, Jewish surnames maintain a significantly higher relative representation; however, in 1941, three years after the introduction of discriminatory racial laws against Jews in Italy by the fascist regime, the relative representation drops to 1.22. From 1951 onwards, with the return of democracy and the beginning of the economic expansion phase, the relative representation of Jewish surnames rises again and in 1961 surpasses the value of 2. Newspapers seem to mirror the changing fortunes of this social group, reflecting societal dynamics over time.

Table 3. RR of surnames groups of Noble, Professors, Jewish and Licei students in newspapers.

Year	**Noble**		**Professors**		**Jewish**		**Licei Students**	
	No	**Yes**	**No**	**Yes**	**No**	**Yes**	**No**	**Yes**
1921	0.54	3.72	0.58	3.42	0.59	4.31	0.53	1.86
	(0.06)	(1.18)	(0.06)	(1.39)	(0.06)	(1.61)	(0.06)	(0.35)
1931	0.63	1.87	0.57	7.27	0.64	4.35	0.56	1.98
	(0.06)	(0.37)	(0.05)	(2.77)	(0.06)	(1.39)	(0.06)	(0.29)
1941	0.64	2.24	0.62	4.21	0.67	1.22	0.56	1.91
	(0.07)	(0.91)	(0.06)	(2.57)	(0.07)	(0.74)	(0.07)	(0.30)
1951	0.55	1.63	0.74	3.28	0.57	1.56	0.53	1.10
	(0.04)	(0.54)	(0.06)	(1.37)	(0.04)	(0.99)	(0.05)	(0.13)
1961	0.65	1.70	0.59	4.44	0.66	2.29	0.62	1.77
	(0.06)	(0.46)	(0.06)	(1.07)	(0.06)	(0.92)	(0.07)	(0.33)
1971	0.80	2.20	0.75	2.70	0.81	2.34	0.77	1.42
	(0.07)	(0.86)	(0.07)	(0.45)	(0.07)	(1.82)	(0.08)	(0.16)
1981	0.78	1.36	0.73	2.64	0.79	1.51	0.77	1.13
	(0.10)	(0.46)	(0.10)	(1.05)	(0.10)	(0.96)	(0.11)	(0.14)
1991	0.55	1.39	0.54	1.45	0.56	0.95	0.52	1.42
	(0.04)	(0.43)	(0.04)	(0.27)	(0.04)	(0.78)	(0.04)	(0.18)
2001	0.55	1.15	0.54	1.42	0.55	3.20	0.52	1.46
	(0.04)	(0.23)	(0.04)	(0.31)	(0.04)	(1.41)	(0.04)	(0.18)
2011	0.55	0.95	0.53	2.08	0.56	1.82	0.53	1.26
	(0.04)	(0.27)	(0.04)	(0.46)	(0.04)	(0.85)	(0.04)	(0.24)

Notes: Standard errors in parentheses. Surnames collected in the subsection's thief, victim, accidents are not included. See Table A.8 in the Appendix for surname groups definitions and statistics.

It must be considered that the section of the newspaper where these surnames appear can influence the results. A valid concern arises regarding the possibility that various articles may not necessarily depict the events of the most influential individuals but rather unrelated incidents not necessarily linked to social status. This concern is particularly pertinent for the section dedicated to Crime News and Court Reporting. Therefore, we excluded surnames associated with articles on crimes, both as victims and perpetrators, from Table 3. Similarly, we excluded all incidents, whether they were road accidents, domestic, or workplace incidents.

For the aforementioned reasons, we anticipate that surnames linked to elite groups will have a lower or in-line relative representation compared to the rest of the population when considering only the subsection covering incidents or thefts in the Crime News and Court Reporting section. This expectation appears to be validated. When we narrow the analysis to surnames collected only in the subsection covering incidents or thefts, the relative representation of surnames belonging to high-status groups decreases significantly, aligning with that of other surnames (see Table A10).

An additional check can be conducted by examining a random group. We expect this group not to show significantly different relative representations from the rest of the population. We extracted the first 500 surnames in alphabetical order from the 1936, 1981, and 2001 records, and observed values exceeding 1 only in 1921 and 1941, while in other years, they were below this threshold and often in line with general trends (see Table A11). Finally, shifting our focus from calculating relative representation in newspapers to that in high schools (see Table A12), socially relevant groups continue to confirm their overrepresentation compared to the rest of the population.

4.2. Mobility of surnames in newspapers

The preceding section has substantiated our hypothesis that many surnames appearing in newspapers are not merely a random extraction from the population but rather contain elements of social relevance/distinction. The way their representation changes over time can thus provide insights into societal dynamics. The more stable the representation of surname groups over time, the lower the mobility, as the same individuals or families tend to be consistently present in newspapers. Conversely, the greater the fluctuation in surname distribution across decades, the more significant the social change and mobility.

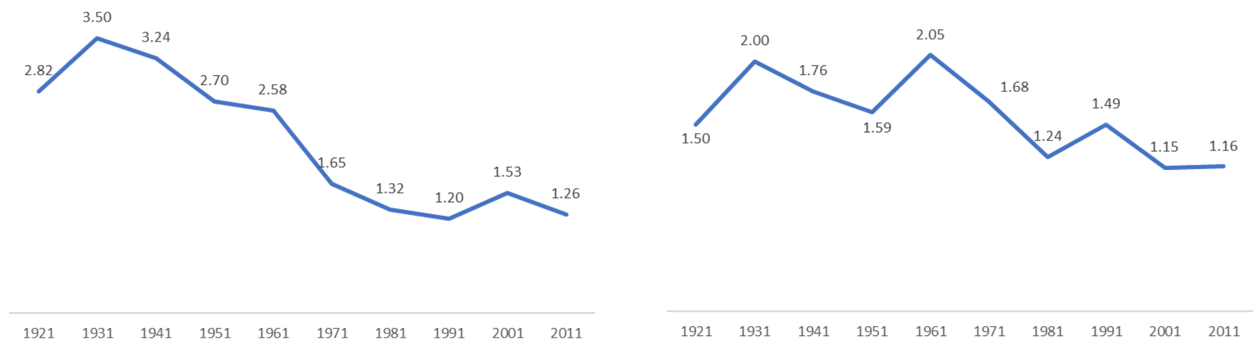
The data at our disposal allow us to differentiate by social group and section. We begin our analysis by focusing on the surnames of university professors and Licei students, as these provide more reliable estimates due to a larger number of observations and a greater variety of surnames. Furthermore, for these groups, we can distinguish by year, with information available for professors teaching in 1921, 1931, and so forth, as well as for students in the same years. For noble surnames and those belonging to the Jewish community, the numerosity and variety of surnames remain stable over time, given the absence of new entries or exits in these groups; for instance, new noble families are not named in the 1900s.

From a broader perspective, examining all newspaper sections except for categories "thief," "victim," and "accidents," reveals changes in the relative representation of surnames of university professors and Licei students who taught and studied in the years 1921-1931-1941-1951, referring to the 1936 registry (Figure 1). Both groups exhibit an RR greater than 1, which would occur if the proportion of surnames in the newspapers mirrored that of the general population. In 1931 and 1941, university professors are present in newspapers over 3 times more than the registry data, while in 1921 and 1951, their presence is more than double their representation in the general population. Over time, the RR tends to decrease, stabilizing around values like 1.2 and 1.3. Conversely, Licei students maintain a constant RR over time, with a slight decrease in the last 40 years considered.

If we narrow the analysis further to only university professors and Licei students in 1931 and 1941, the years closest to the 1936 registry, the RR of professors in those years becomes even more pronounced (Figure 2). In newspapers from 1931 and 1941, they are present almost 5 times more than the registry data, but over time, their RR decreases, converging towards the mean (RR=1) from the 1980s. Licei students, on the other hand, have a lower RR, just over 2 times their representation in the registry data in 1931 and 1.76 in 1941. A regression towards the mean is also observed for them, though it is slower and less pronounced.

Figure 1. RR of university professors and Licei students who taught and studied in 1921-1931-1941-1951.

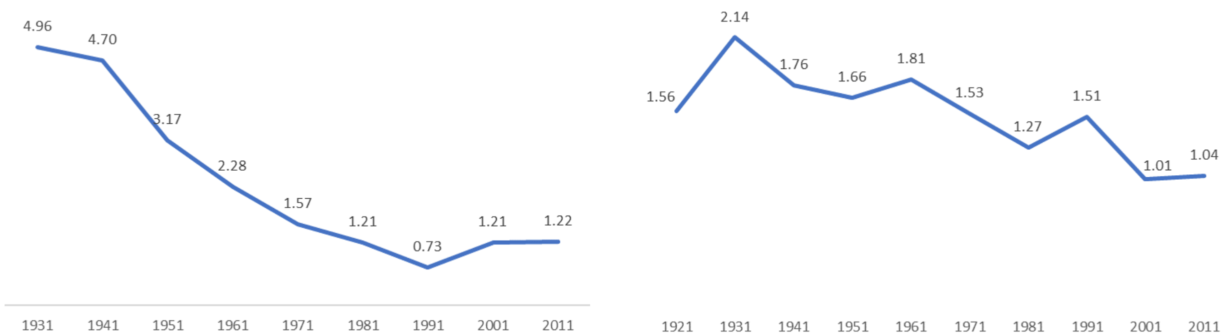
University Professors 1921-1931-1941-1951 *Licei students 1921-1931-1941-1951*



Notes: Surnames collected in the subsection's thief, victim, accidents are not included. See Table A.8 in the Appendix for surname groups definitions and statistics.

Figure 2. RR of university professors and Licei students who taught and studied in 1931 and 1941.

University Professors 1931-1941 *Licei students 1931-1941*



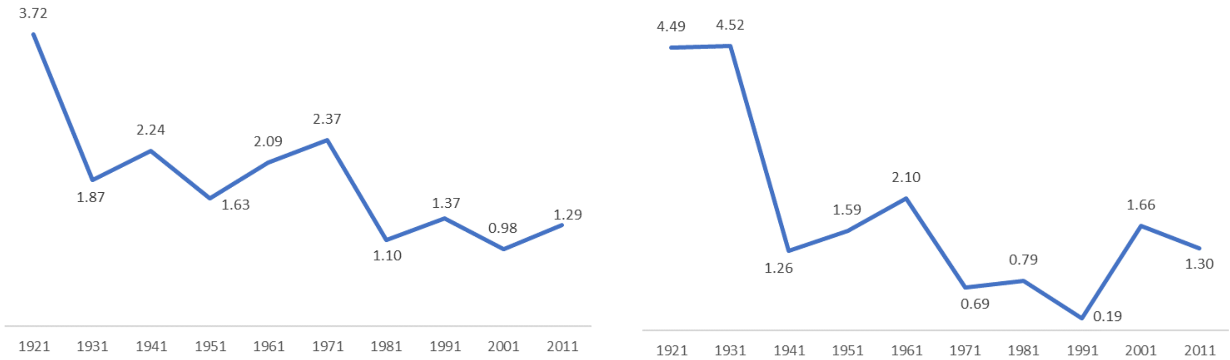
Notes: Surnames collected in the subsection's thief, victim, accidents are not included. See Table A.8 in the Appendix for surname groups definitions and statistics.

As previously mentioned, the analysis of noble surnames and those belonging to the Jewish community is constrained by the stability of these data over time, marked by the absence of new entries or exits within these surname groups. The only distinction feasible is to consider, for years after '51, when the reference registry transitions to that of 1981 and 2001, only those surnames present in the 1936 registry. Consequently, a noble surname is considered in 1961 only if it was also present in the 1936 registry.

The results confirm a higher prevalence of noble surnames in newspapers compared to their representation in the general population in the early 1900s, particularly in 1921 (Figure 3). This trend remains robust until the 1970s, converging towards the mean from the 1980s onward. Surnames of Jewish origin exhibit a higher relative representation in the pre-World War II years, sharply declining in 1941 following the introduction of racial laws in Italy and the country's alignment with Nazi Germany. Their RR experiences a slight increase in 1951 and 1961 but remains significantly lower compared to the pre-change in policy towards them by the Italian fascist government.

Figure 3. RR of noble and Jewish community surnames

Noble *Jewish*

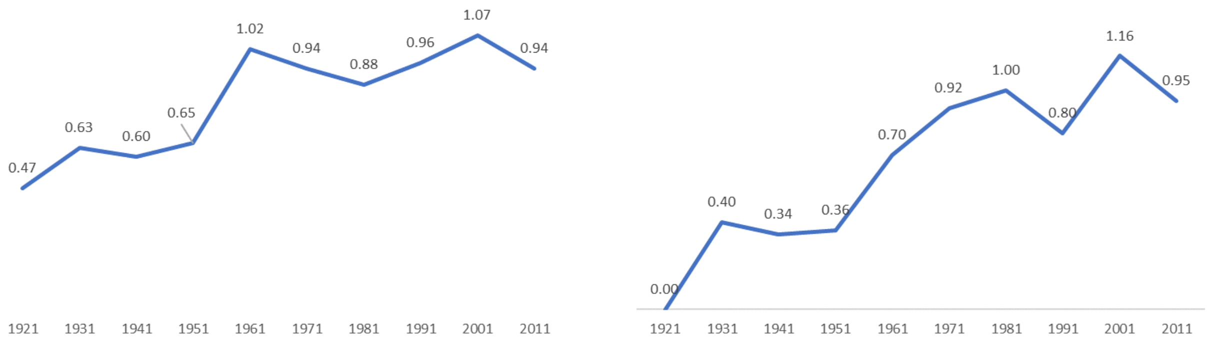


Notes: Surnames collected in the subsection's thief, victim, accidents are not included. See Table A.8 in the Appendix for surname groups definitions and statistics.

For the most influential social classes, it is anticipated that a higher RR will decrease over time. Similarly, for those poorly represented or even absent in newspapers in a given year, a convergence towards the mean is expected in the medium to long term. This phenomenon is particularly evident for surnames with a low RR but a widespread presence in the general population. Figure 4 provides a clear illustration of this concept: among the 200 most common surnames in 1936, we select the 140 with low RR in newspapers in 1921 (below the 60th percentile of the median). Overall, their RR is 0.47 in 1921. An increase is observed in the following years, stabilizing around the threshold of equal distribution of 1 starting from the 1960s. Focusing the analysis on the 53 surnames without mentions in newspapers in 1921 reveals a similar trend. There is a consistent regression towards the mean converging to 1 by 1981, with a particularly pronounced increase in the 1960s and 1970s.

Figure 4. RR of common and underrepresented surnames in 1921 newspapers

Below the 60th percentile of the median *No representation*



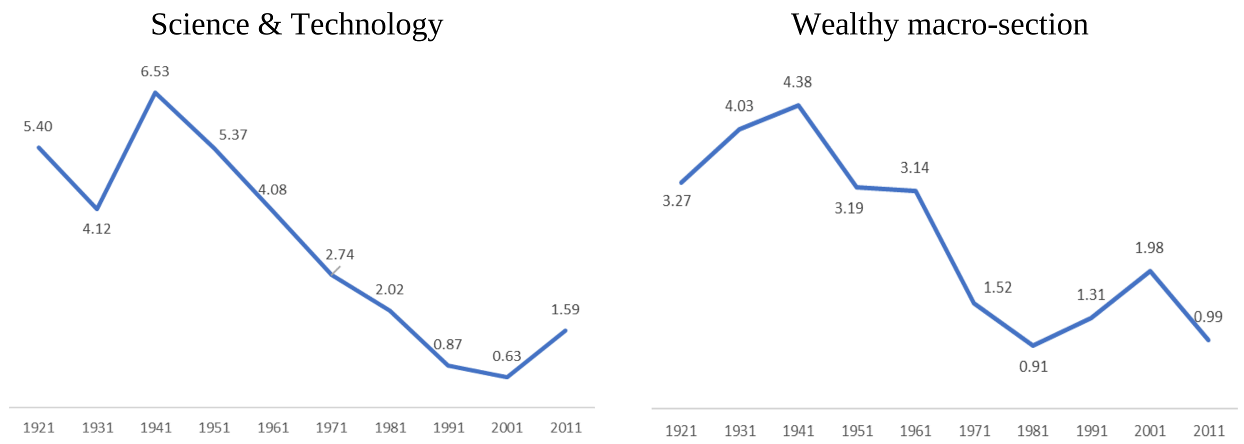
Notes: Surnames collected in the subsection's thief, victim, accidents are not included. See Table A.8 in the Appendix for surname groups definitions and statistics.

The analysis of data across newspaper sections reveals additional aspects of interest. For instance, we can investigate the mobility of surnames associated with university professors who taught in the years

1921-1931-1941-1951, specifically focusing on the section directly relevant to their field, namely "Science and Technology" (Figure 5). Their relative representation is significantly high during the years directly relevant to them (from 1921 to 1951), surpassing that observed when considering all newspaper sections. This elevated representation persists in the subsequent years before regressing towards the mean from the 1990s onward.

As one would expect, professions associated with privilege, such as university professors, are anticipated to correlate with a better economic status. Therefore, it is not surprising to observe that professors from 1921-1931-1941-1951 exhibit a high Relative Representation in the macro-section labeled "wealthy"[5] during those respective years (Figure 5). Although their representation diminishes over time in this macro-section, it consistently remains above 1.

Figure 5. RR of university professors who taught in 1921-1931-1941-1951 in Science & Technology section and Wealthy macro-section.

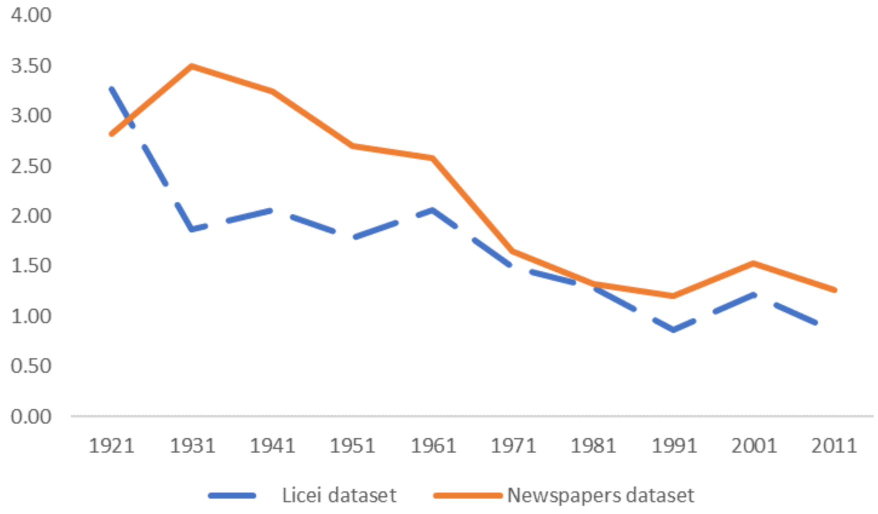


Notes: See Table A.8 in the Appendix for surname groups definitions and statistics.

In the final stage, a comparison is drawn between the results obtained from newspapers and the dataset of high schools. We consider the surnames of professors from the years 1921-1931-1941-1951 and, in addition to studying their RR in newspapers, we also analyze it in the Licei dataset. The expected outcome remains consistent; namely, surnames associated with professors are anticipated to have a higher representation in Licei compared to their presence in the general population (Figure 6). While the two curves do not directly overlap, they exhibit a relatively similar trend. In newspapers, there is a notable emphasis on the RR of the elite group, particularly in 1931 and 1941.

[5] The "wealthy" macrosection refers to the combination of the Advertising, Business, Obituaries sections, and the subsection Judge-Lawyer-Police

Figure 6. RR of university professors who taught in 1921-1931-1941-1951 in Newspaper and Licei dataset



Notes: Surnames collected in the subsection's thief, victim, accidents are not included. See Table A.8 in the Appendix for surname groups definitions and statistics.

5. Robustness check

Given the nature of our data, two robustness checks must be carried out. The first concerns the use of surnames; in fact, it is not possible to state with certainty that a surname observed today is part of the direct descendants of the same surname found in previous decades. One method to strengthen these pseudo-connections is to limit the investigation to the rarest surnames. We repeat the analyses previously performed by excluding the 200 most common surnames in the 1936 registry and the 600 most common surnames in the 1981 and 2001 registries. This allows us to have a maximum of 20 families with the same surname in 1936 and a maximum of 50 individuals with the same surname in 1981 and 2001 (consistent with an average family size of 2.5 people). It is also important to consider that the vast majority of surnames belong to only one family in 1936 (62.17%, see Table A5 in the appendix). If we add surnames belonging to a maximum of 2 or 3 families, we cover almost 80% of the total population. The same applies to 1981 and 2001; considering surnames belonging to a maximum of 5 individuals covers 80% of the entire population. Table 4 reproduces the setup of Table 3, showing an RR of rare surnames among elite groups in line with or even higher than that observed for all surnames. Similar trends, but with higher RRs for rare surnames, are also confirmed for other analyses (see Figure A1 in the appendix, reproducing Figure 2)

Table 4. RR of rare surnames groups of Noble, Professors, Jewish and Licei students in newspapers.

Year	**Noble**		**Professors**		**Jewish**		**Licei Students**	
	No	**Yes**	**No**	**Yes**	**No**	**Yes**	**No**	**Yes**
1921	0.71	4.36	0.73	4.31	0.77	4.49	0.71	2.34
	(0.08)	(1.45)	(0.08)	(1.85)	(0.08)	(1.67)	(0.08)	(0.51)
1931	0.82	2.14	0.75	9.60	0.83	4.52	0.75	2.39
	(0.08)	(0.45)	(0.06)	(3.82)	(0.08)	(1.45)	(0.08)	(0.40)
1941	0.83	2.60	0.81	5.31	0.87	1.26	0.75	2.31
	(0.09)	(1.12)	(0.08)	(3.42)	(0.10)	(0.78)	(0.10)	(0.40)
1951	0.72	1.82	0.70	3.85	0.74	1.59	0.70	1.29
	(0.06)	(0.67)	(0.06)	(1.63)	(0.06)	(1.04)	(0.06)	(0.19)
1961	0.68	2.15	0.62	5.88	0.68	2.43	0.65	2.32
	(0.07)	(0.72)	(0.07)	(1.54)	(0.07)	(0.99)	(0.07)	(0.58)
1971	0.84	2.90	0.80	3.37	0.85	2.47	0.81	1.78
	(0.08)	(1.36)	(0.08)	(0.65)	(0.08)	(1.95)	(0.08)	(0.26)
1981	0.82	1.46	0.77	3.34	0.82	1.58	0.81	1.22
	(0.11)	(0.72)	(0.11)	(1.55)	(0.11)	(1.02)	(0.11)	(0.23)
1991	0.56	1.50	0.55	1.61	0.56	0.96	0.54	1.61
	(0.04)	(0.65)	(0.04)	(0.40)	(0.04)	(0.85)	(0.04)	(0.29)
2001	0.56	1.08	0.55	1.49	0.55	3.35	0.53	1.65
	(0.04)	(0.33)	(0.04)	(0.46)	(0.04)	(1.53)	(0.04)	(0.28)
2011	0.56	1.06	0.54	2.61	0.56	1.94	0.55	1.39
	(0.04)	(0.42)	(0.04)	(0.69)	(0.04)	(0.92)	(0.04)	(0.39)

Notes: Standard errors in parentheses. Surnames collected in the subsection's thief, victim, accidents are not included. See Table A.8 in the Appendix for surname groups definitions and statistics.

The second robustness check concerns newspapers. Throughout the 20th century, just as society has changed, so have newspapers, their format, the number of pages, the events they choose to highlight, and consequently, their informational content. Table 1 and especially Table A1 in the appendix highlight this aspect. For instance, there are significant variations in the "Public Events & Civil Society" section, which has gained increasing space from the 1970s and even more so in the 2000s. To ensure robust estimates, information provided in local newspapers should ideally remain relatively constant. Otherwise, observed changes may reflect a decrease (or increase) in the space given to certain articles relative to others. For this reason, we conduct a verification of the results that take these considerations into account by reweighting surnames.

If the incidence of a section on the total observations in a particular year exceeds 22.75%, then that citation is weighted 0.5 instead of 1. If it exceeds 40%, it is weighted 0.33. This approach does not completely shield us from the highlighted problem; some sections are abandoned over time and cannot be recovered (see Charity from 1971 onwards). However, it allows to obtain a more stable dynamic. Table A13 and Figure A2 reiterate the analyses in Table 1 and Figure 2, but with the new weighting. The results are almost entirely overlapping with those obtained previously.

Finally, it is worth considering that changes in newspapers are closely linked to changes in society. Therefore, if editors decide to give more space to certain articles over others, it may be a result of societal shift, with readers being more interested in specific events and those mentioned in these articles. Thus, the fundamental hypothesis of this article is not lost, namely that those who appear more frequently in newspapers likely have a higher influence in society.

6. Conclusions

In this study, we propose the use of surnames in local newspapers as a novel data source to explore social mobility of elites and family generations. The key idea is that individuals frequently cited in newspapers hold social importance, and we aim to investigate whether this higher visibility persists over time within the same families.

We first verified that surnames in newspapers are not randomly extracted from the population but reflect traits of social relevance. Subsequently, we analyzed the mobility of surnames in newspapers over time. Surnames belonging to privileged groups exhibit a higher representation compared to other social groups, and this higher RR seems to be transmitted to future generations in a differentiated manner depending on the considered high-status group. In the case of surnames of university professors and Licei students, their representation tends to converge towards the mean, particularly from the '80s onwards. The same holds true for noble surnames, while for surnames belonging to the Jewish community, the pivotal moment of change is situated in '41.

We compared the results obtained from newspapers with a more common indicator in the study of social status, namely education, noting a relatively similar trend between the two information sources. However, newspapers appear to overstate the representation of relevant social groups compared to education. We also performed controls for rare surnames and reweighted the newspaper dataset considering changes in newspaper structure over time. Both controls do not show significantly different results from our main analysis.

It would be interesting to replicate this exercise with other local contexts to verify if the outcomes are confirmed. According to Acciari et al. (2022), northern Italian territories are characterized by substantial economic mobility; therefore, a comparison with a local reality in different parts of Italy could be very interesting to separate the influence of common historical changes from that of local factors. Future studies could replicate this analysis also in contexts outside Italy and for different elite groups, including a third important information source as a comparison, such as incomes or occupational status. Moreover, it would be impactful to repeat this approach with digital newspaper archives and using automated data collection techniques to reduce time and bias in data collection.

In conclusion, we suggest that this type of analysis could contribute to identifying different mobility patterns at the local level and represent a useful alternative when established data sources, especially income data, are not available.

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Appendix

Table A1. Sections incidence on the total collected surnames per year.

Section	1921	1931	1941	1951	1961	1971	1981	1991	2001	2011
Advertising	15.30%	19.21%	11.91%	18.53%	24.62%	8.37%	6.49%	10.60%	5.87%	1.54%
Business	8.84%	10.82%	5.94%	6.98%	6.43%	10.84%	9.77%	2.06%	7.12%	3.55%
Charity	15.47%	28.96%	18.11%	10.67%	11.77%	5.70%	2.57%	0.00%	0.00%	0.00%
Culture & Public Events	9.14%	6.12%	17.98%	16.35%	14.55%	25.93%	26.88%	27.93%	42.15%	45.50%
Local News	22.80%	8.53%	13.55%	22.17%	3.29%	10.82%	9.24%	4.94%	4.84%	4.61%
Obituaries	2.37%	1.25%	12.60%	6.02%	2.67%	7.59%	5.36%	4.41%	5.18%	4.25%
Local Politics	22.89%	13.55%	15.62%	14.21%	21.89%	21.25%	28.36%	40.16%	26.13%	32.16%
Religion	0.00%	0.93%	0.66%	1.43%	1.60%	1.08%	2.22%	2.58%	1.57%	1.44%
Science & Technology	3.19%	10.62%	3.64%	3.64%	13.18%	8.42%	9.10%	7.33%	7.13%	6.95%
Total	**100%**	**100%**	**100%**	**100%**	**100%**	**100%**	**100%**	**100%**	**100%**	**100%**

Note: The Crime News & Court Reporting section is a combination of the subsections' thief, victim, accidents, and judges-lawyers-police.

Table A2. Statistics on surnames in the Newspaper Dataset

	1921	1931	1941	1951	1961
Unit of measurement	Individual	Individual	Individual	Individual	Individual
Number of different surnames	802	1,033	1,021	1,226	1,115
Total observations	2,320	3,446	3,048	4,068	3,066
	1971	**1981**	**1991**	**2001**	**2011**
Unit of measurement	Individual	Individual	Individual	Individual	Individual
Number of different surnames	1,458	1,213	1,224	1,792	1,625
Total observations	3,717	2,835	3,058	5,537	4,991

Table A3. Rarity of surnames in the Newspaper Dataset

	Number of observations of the same surname	Surnames per frequency class	frequency	Cumulative frequency		Number of observations of the same surname	Surnames per frequency class	frequency	Cumulative frequency
	1	467	58.23%	58.23%		*1*	886	60.77%	60.77%
	2 to 3	197	24.56%	82.79%		*2 to 3*	374	25.65%	86.42%
1921	*4 to 10*	113	14.09%	96.88%	***1971***	*4 to 10*	166	11.39%	97.81%
	11 to 30	23	2.87%	99.75%		*11 to 30*	29	1.99%	99.79%
	31 and above	2	0.25%	100.00%		*31 and above*	3	0.21%	100.00%
	1	549	53.15%	53.15%		*1*	756	62.32%	62.32%
	2 to 3	224	21.68%	74.83%		*2 to 3*	281	23.17%	85.49%
1931	*4 to 10*	199	19.26%	94.09%	***1981***	*4 to 10*	153	12.61%	98.10%
	11 to 30	59	5.71%	99.81%		*11 to 30*	22	1.81%	99.92%
	31 and above	2	0.19%	100.00%		*31 and above*	1	0.08%	100.00%
	1	611	59.84%	59.84%		*1*	728	59.48%	59.48%
	2 to 3	230	22.53%	82.37%		*2 to 3*	292	23.86%	83.33%
1941	*4 to 10*	142	13.91%	96.28%	***1991***	*4 to 10*	169	13.81%	97.14%
	11 to 30	32	3.13%	99.41%		*11 to 30*	34	2.78%	99.92%
	31 and above	6	0.59%	100.00%		*31 and above*	1	0.08%	100.00%
	1	738	60.20%	60.20%		*1*	971	54.19%	54.19%
	2 to 3	274	22.35%	82.54%		*2 to 3*	455	25.39%	79.58%
1951	*4 to 10*	160	13.05%	95.60%	***2001***	*4 to 10*	282	15.74%	95.31%
	11 to 30	49	4.00%	99.59%		*11 to 30*	77	4.30%	99.61%
	31 and above	5	0.41%	100.00%		*31 and above*	7	0.39%	100.00%
	1	656	58.83%	58.83%		*1*	931	57.29%	57.29%
	2 to 3	259	23.23%	82.06%		*2 to 3*	362	22.28%	79.57%
1961	*4 to 10*	159	14.26%	96.32%	***2011***	*4 to 10*	245	15.08%	94.65%
	11 to 30	36	3.23%	99.55%		*11 to 30*	79	4.86%	99.51%
	31 and above	5	0.45%	100.00%		*31 and above*	8	0.49%	100.00%

Table A4. Statistics on surnames in registries

	1936	**1981**	**2001**
Unit of measurement	Family	Individual	Individual
Number of different surnames	4,407	16,858	24,710
Observations	21,115	180,459	178,017
Cumulative frequency of the 10 most common surnames	7.58%	5.73%	4.57%
Cumulative frequency of the 100 most common surnames	33.65%	24.46%	19.46%
Cumulative frequency of the 200 most common surnames	48.30%	35.63%	28.51%

Table A5. Rarity of surnames in registries

	Number of observations of the same surname	**Surnames per frequency class**	**frequency**	**Cumulative frequency**
1936	*1*	2,740	62.17%	62.17%
	2	498	11.30%	73.47%
	3	239	5.42%	78.90%
	4 to 5	256	5.81%	84.71%
	6 to 10	260	5.90%	90.61%
	11 to 20	184	4.18%	94.78%
	21 to 50	158	3.59%	98.37%
	51 to 100	61	1.38%	99.75%
	101 to 200	9	0.20%	99.95%
	201 to 500	2	0.05%	100.00%
	501 to 1000	0	0.00%	100.00%
	1000 and above	0	0.00%	100.00%
1981	*1*	6,789	40.27%	40.27%
	2	2,290	13.58%	53.86%
	3	1,890	11.21%	65.07%
	4 to 5	1,785	10.59%	75.66%
	6 to 10	1,686	10.00%	85.66%
	11 to 20	1,014	6.01%	91.67%
	21 to 50	735	4.36%	96.03%
	51 to 100	331	1.96%	98.00%
	101 to 200	190	1.13%	99.12%
	201 to 500	125	0.74%	99.86%
	501 to 1000	20	0.12%	99.98%
	1000 and above	3	0.02%	100.00%
2001	*1*	11,289	45.69%	45.69%
	2	3,513	14.22%	59.90%
	3	2,537	10.27%	70.17%
	4 to 5	2,578	10.43%	80.60%
	6 to 10	2,199	8.90%	89.50%
	11 to 20	1,174	4.75%	94.25%
	21 to 50	834	3.38%	97.63%
	51 to 100	318	1.29%	98.92%
	101 to 200	164	0.66%	99.58%
	201 to 500	92	0.37%	99.95%
	501 to 1000	10	0.04%	99.99%
	1000 and above	2	0.01%	100.00%

Table A6. Statistics on surnames in the Licei Dataset

	1921	1931	1941	1951	1961
Unit of measurement	Individual	Individual	Individual	Individual	Individual
Number of different surnames	440	450	615	587	718
Total observations	628	614	885	838	1,069
	1971	**1981**	**1991**	**2001**	**2011**
Unit of measurement	Individual	Individual	Individual	Individual	Individual
Number of different surnames	1,208	992	1,061	999	926
Total observations	2,164	1,510	1,619	1,407	1,224

Table A7. Rarity of surnames in the Licei Dataset

	Number of observations of the same surname	Surnames per frequency class		Number of observations of the same surname	Surnames per frequency class
1921	*1*	314	***1971***	*1*	820
	2 to 3	115		*2 to 3*	277
	4 to 9	10		*4 to 9*	103
	10 and above	1		*10 and above*	8
1931	*1*	337	***1981***	*1*	725
	2 to 3	104		*2 to 3*	216
	4 to 9	8		*4 to 9*	58
	10 and above	1		*10 and above*	3
1941	*1*	454	***1991***	*1*	796
	2 to 3	141		*2 to 3*	200
	4 to 9	19		*4 to 9*	63
	10 and above	1		*10 and above*	2
1951	*1*	435	***2001***	*1*	772
	2 to 3	134		*2 to 3*	187
	4 to 9	9		*4 to 9*	38
	10 and above	1		*10 and above*	2
1961	*1*	525	***2011***	*1*	740
	2 to 3	161		*2 to 3*	165
	4 to 9	30		*4 to 9*	19
	10 and above	2		*10 and above*	2

Table A8. Surname groups definition, source, and statistics.

Group	Definition	Source	Statistics	1936	1981	2001
Noble	Surnames of Modenese noble families	Giacomo Pietramellara, *Blasonario generale italiano, ossia Descrizione degli stemmi delle famiglie nobili e titolate d'Italia. Dispensa 6: del modenese.* Direzione, Roma 1902.	*Total number of surnames*	299	299	299
			Cumulative frequency of the 10 most common surnames	6.30%	4.57%	3.58%
			Total cumulative frequency	11.58%	9.04%	7.31%
Professor	Surnames of professors at the University of Modena	Academic yearbooks of the University of Modena. Years 1921, 1931, 1941, 1951, 1961, 1971, 1981, 2003	*Total number of surnames*	1544	1544	1544
			Cumulative frequency of the 10 most common surnames	6.73%	5.40%	4.31%
			Total cumulative frequency	18.90%	29.72%	26.67%
Jew	Surnames of Modenese Jewish families	Franco Bonilauri, Vincenza Maugeri, *Le comunità ebraiche a Modena e a Carpi: dal Medioevo all'età contemporanea*, Giuntina, 1999	*Total number of surnames*	42	42	42
			Cumulative frequency of the 10 most common surnames	0.75%	0.43%	0.37%
			Total cumulative frequency	0.90%	0.49%	0.42%
Common and underrepresented surnames	Common surnames in 1936 registry with low representation in 1921 newspaper (below the 60th percentile of the median)	Our elaboration on 1936 registry	*Total number of surnames*	140	140	140
			Cumulative frequency of the 10 most common surnames	4.64%	3.58%	2.86%
			Total cumulative frequency	29.71%	20.82%	16.47%
Random	First 500 surnames in alphabetical order of 1936-1981-2001 registries	Our elaboration	*Total number of surnames*	500	500	500
			Cumulative frequency of the 10 most common surnames	3.70%	0.61%	0.26%
			Total cumulative frequency	12.22%	1.81%	0.88%

Table A9. Correlation between the RR in newspapers of the 200 most common surnames in general population and their frequency distribution in general population.

	RR 1921	**RR 1931**	**RR 1941**	**RR 1951**	**RR 1961**	**RR 1971**	**RR 1981**	**RR 1991**	**RR 2001**	**RR 2011**
Frequency distribution registry 1936	-0.00	-0.10	-0.02	-0.04						
Frequency distribution registry 1981					-0.01	-0.05	-0.04			
Frequency distribution registry 2001								-0.07	-0.02	-0.07

Note: Frequency distributions in general population refer exclusively to surnames present in newspapers during the reference year. Therefore, the -0.23 value for 1921 is the outcome of the correlation between the RR of surnames present in 1921 newspapers and their frequency distribution in the 1936 registry.

Table A10. RR of surnames groups of Noble, Professors, Jewish and Licei students in newspapers subsection's thief and accidents

Year	Noble		Professors		Jewish		Licei students	
	No	Yes	No	Yes	No	Yes	No	Yes
1921	0.49	0.89	0.49	1.02	0.50	0.00	0.49	0.71
1931	0.89	4.63	1.00	0.28	1.00	0.00	0.94	1.58
1941	0.53	1.12	0.55	0.44	0.55	0.00	0.47	1.22
1951	0.62	0.78	0.61	1.37	0.62	0.00	0.57	1.10
1961	0.38	0.29	0.37	0.58	0.38	0.17	0.37	0.55
1971	0.83	0.75	0.82	1.06	0.83	0.00	0.84	0.66
1981	1.39	0.58	1.40	0.84	1.38	3.21	1.43	0.39
1991	0.97	1.37	0.98	0.65	0.98	0.00	0.97	1.08
2001	0.89	0.21	0.89	0.24	0.88	0.00	0.90	0.31
2011	0.64	0.08	0.64	0.16	0.63	0.45	0.65	0.18

Notes: See Table A.8 in the Appendix for surname groups definitions and statistics.

Table A11. RR of surnames of random group in newspapers (first 500 surnames in alphabetical order)

Year	Noble	
	No	Yes
1921	0.56	1.14
1931	0.65	0.71
1941	0.63	1.05
1951	0.57	0.65
1961	0.70	0.65
1971	0.86	0.58
1981	0.83	0.85
1991	0.59	0.02
2001	0.58	0.33
2011	0.58	0.35

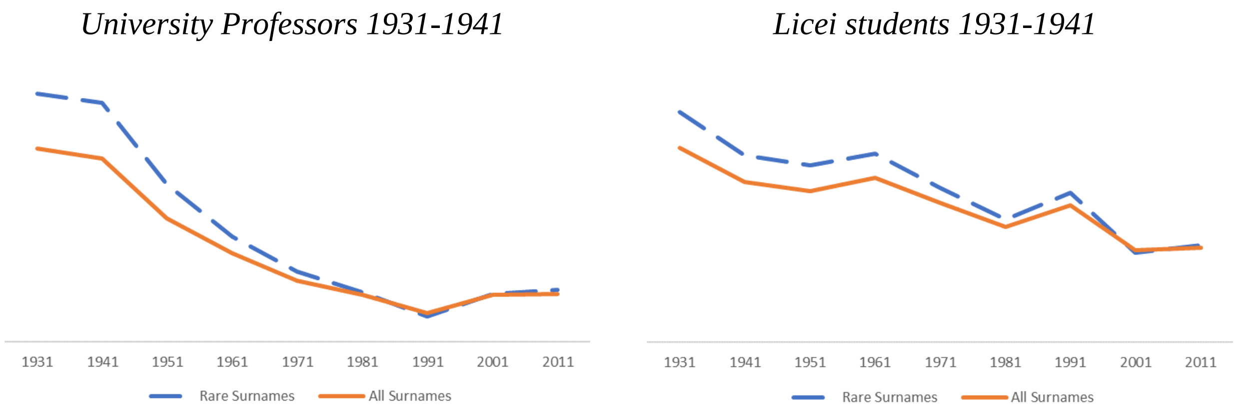
Notes: Surnames collected in the subsection's thief, victim, accidents are not included. See Table A.8 in the Appendix for surname groups definitions and statistics.

Table A12. RR of surnames groups of Noble, Professors, Jewish in Licei dataset.

Year	Noble		Professors		Jewish	
	No	**Yes**	**No**	**Yes**	**No**	**Yes**
1921	0.71	1.77	0.68	3.32	0.73	3.21
	(0.07)	(0.39)	(0.07)	(1.04)	(0.07)	(1.15)
1931	0.84	1.29	0.84	1.68	0.84	2.55
	(0.07)	(0.31)	(0.07)	(0.43)	(0.07)	(0.90)
1941	0.88	1.56	0.87	2.27	0.89	1.52
	(0.06)	(0.37)	(0.06)	(0.59)	(0.06)	(0.68)
1951	0.66	1.31	0.66	2.18	0.67	1.52
	(0.06)	(0.33)	(0.05)	(1.11)	(0.05)	(1.12)
1961	0.72	1.25	0.69	2.52	0.72	1.71
	(0.29)	(0.06)	(0.06)	(0.75)	(0.06)	(0.83)
1971	0.66	0.79	0.63	1.83	0.66	2.58
	(0.04)	(0.14)	(0.04)	(0.36)	(0.04)	(1.22)
1981	0.70	1.46	0.64	2.84	0.70	4.65
	(0.05)	(0.39)	(0.05)	(0.42)	(0.05)	(1.90)
1991	0.56	2.05	0.54	2.10	0.57	1.29
	(0.04)	(0.96)	(0.04)	(0.47)	(0.04)	(0.79)
2001	0.59	1.57	0.59	1.42	0.60	0.66
	(0.04)	(0.81)	(0.04)	(0.37)	(0.04)	(0.40)
2011	0.58	1.38	0.57	1.44	0.59	1.10
	(0.04)	(0.49)	(0.04)	(0.34)	(0.04)	(0.65)

Notes: Standard errors in parentheses . Surnames collected in the subsection's thief, victim, accidents are not included. See Table A.8 in the Appendix for surname groups definitions and statistics.

Figure A1. RR of rare and all surnames of university professors and Licei students who taught and studied in 1931 and 1941.



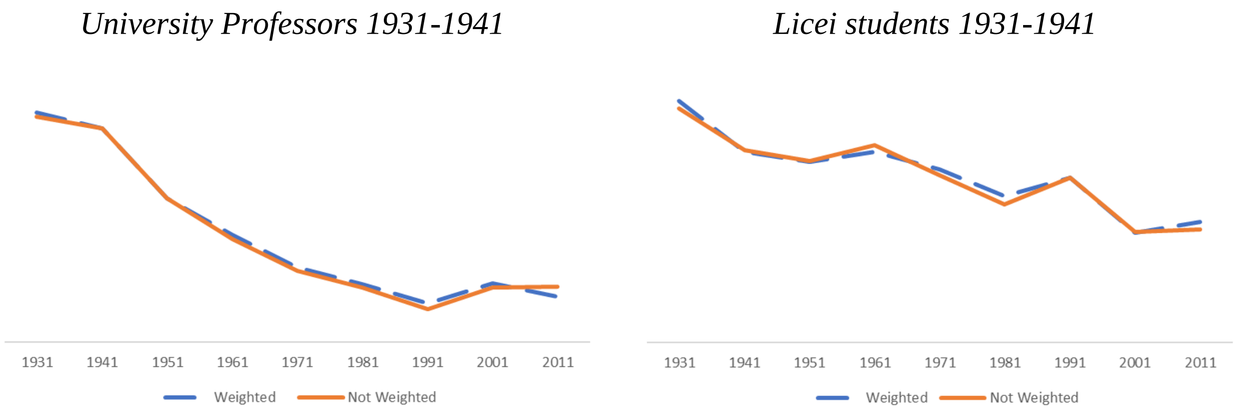
Notes: Surnames collected in the subsection's thief, victim, accidents are not included. See Table A.8 in the Appendix for surname groups definitions and statistics.

Table A13. RR of weighted surnames groups of Noble, Professors, Jewish and Licei students in newspapers

Year	Noble		Professors		Jewish		Licei students	
	No	**Yes**	**No**	**Yes**	**No**	**Yes**	**No**	**Yes**
1921	0.54	3.27	0.54	3.65	0.59	4.00	0.53	1.83
1931	0.64	1.85	0.59	7.70	0.66	4.17	0.58	2.02
1941	0.64	2.24	0.62	4.21	0.67	1.22	0.56	1.89
1951	0.55	1.63	0.55	3.32	0.57	1.56	0.53	1.10
1961	0.67	1.65	0.61	4.48	0.68	2.46	0.64	1.81
1971	0.80	2.31	0.76	2.86	0.81	2.51	0.77	1.47
1981	0.83	1.47	0.77	3.04	0.83	1.82	0.82	1.15
1991	0.60	1.72	0.59	1.84	0.61	1.46	0.58	1.28
2001	0.54	1.11	0.53	1.56	0.54	3.11	0.51	1.45
2011	0.55	0.63	0.51	2.88	0.55	2.35	0.53	1.13

Notes: See Table A.8 in the Appendix for surname groups definitions and statistics.

Figure A4. RR of weighted surnames of university professors and Licei students who taught and studied in 1931 and 1941



Notes: Surnames collected in the subsection's thief, victim, accidents are not included. See Table A.8 in the Appendix for surname groups definitions and statistics.

Chapter 3:
The biased reaction to changes in family-related public expenditure: How generosity and universalism influence fertility *

Abstract

This study examines the relationship between fertility and social policies across countries within the European Union. Based on 2004-2020 EU-SILC data, the research investigates how relevant changes in family/children allowances influence the likelihood of new births in the short term. Based on the social investment hypothesis and a general expansion of family policies in the EU since 2004, we investigate if and how increased family support contributes to birth events within families. The analysis assesses asymmetric reactions in fertility to changes in family-related social benefits, specifically looking at changes in their generosity and universalism levels from one year to another. Findings indicate that enhancing the generosity of cash benefits is positively correlated with an increase in the likelihood of having a child, while the opposite occurs in case of reduction of the transfers universalism. Results of the heterogeneity analysis also reveal that the reaction to changes in public spending is not the same across the population but differs according to characteristics of mothers.

Keywords: fertility; family allowances; public expenditure; social policies; EU countries.

JEL classification codes: I38, J13, J18.

* This chapter is co-authored with Giovanni Gallo (University of Modena and Reggio Emilia) and Stephan Köppe (University College Dublin).

1. Introduction

In several European countries, there is growing concern about demographic trends resulting from population ageing. This has been driven over time by a combination of increased life expectancy and decreasing fertility rates. In particular, total fertility rate (TFR) began to decline very rapidly from the late 1960s (Caldwell and Schindlmayr, 2003). As early as 1977, the average TFR of European Union countries (EU27) were below the replacement level. The short-term recovery that occurred since 2000, which had made scholars wonder if the era of "Lowest-Low Fertility" was over (Goldstein et al. 2009), also seems to have lost its momentum after a decade. Today numerous countries present birth rates in line with levels recorded in the mid-1990s. [12]

In such a context, it is relevant to ask how policies – and in particular family policies – can somehow reverse these trends by influencing fertility levels. Indeed, on the one hand the effect of fertility on the welfare state is widely recognized, with significant implications for the sustainability of the welfare system, social cohesion, economic dynamics, and intergenerational equity. On the other hand, the role of policy interventions on the TFR is less clear (Moffitt, 1998; Gauthier 2007). According to neoclassical economic theory of fertility, the decision to have children is considered a rational choice based on the maximization of economic utility. This decision depends on economic costs and benefits of parenthood, considering the income constraint and individual preferences for children over other goods (Becker, 1991). It implies that any reduction in the costs of children, such as through public subsidies for child-related services, or any increase in income, such as cash transfers, should increase the demand for children. This economic model has exerted a strong effect in the literature and underlies the assumed relationship between policies and fertility. However, it is based on a simplified model with restrictive key assumptions, which have implications for the relationship between policies and fertility that may explain some of the unexpected or inconsistent results that have emerged in empirical research (Gauthier, 2007). The basic theoretical model also underappreciates the gendered nature of fertility decisions. The social investment literature has consistently highlighted the mother penalty in the labor market (Aisenbrey et al. 2009), where career-oriented women are faced with a choice between children or a career. Labor market-oriented family policies, pioneered in the Nordic countries, such as childcare support and parental leave (Ferrarini, 2006) were designed to increase female employment and reduce motherhood penalties. Yet, empirical evidence on the extent to which fertility is affected by family-friendly labor market policies, such as maternity leave and childcare, is mixed and inconclusive: some scholars claim that work-related benefits have a small positive influence on fertility, while others find no evidence of an effect (Kalwij, 2010). Findings on the effect

[12] Available OECD data on TFR trend by country at: https://data.oecd.org/pop/fertility-rates.htm.

of family or child cash benefits on fertility instead appeared to be more aligned with the theory, suggesting a positive and significant (albeit small) effect of cash transfers on fertility. More recent contributions, however, have found mixed evidence, with some scholars continuing to emphasize a positive effect between family allowances and fertility, while others find no significant effects. Part of the mixed evidence is due to how fertility is operationalized. Results differ if considering total cohort fertility, childless women, or large families, as well as families with different income levels and working status. The present study builds on this literature.

We investigate the relationship between birth events and levels of expenditure on family policies, focusing on family/children-related allowances, through a comparative analysis of European Union countries. Specifically, we examine whether and to what extent relevant changes in the public provision of family-child-related allowances affect the likelihood of reporting a birth event in the short term. To be clear, we are not concerned with fertility intentions or the effect of family policies on employment outcomes (see Finch and Bradshaw, 2021).

Drawing on the social investment literature (Finch and Bradshaw, 2021; Jenson and Mahon, 2022), our focus centers on the generosity and universalism of social spending, exploring the hypotheses that increasing cash transfers and/or broadening the number of recipient families have a positive effect on the likelihood of giving birth, while the opposite occurs in case of public expenditure retrenchment. As a large behavioral literature shows that individuals are prone to a negativity bias when dealing with an economic or information shock (Kahneman and Tversky, 1979; Baumeister et al., 2001; Eil and Rao, 2011; De Neve, 2018), this paper also aims to assess the presence of an asymmetric reaction in terms of birth events to changes in generosity and universalism of family-related social transfers. In line with the behavioral literature, if individuals operate in a condition of perfect information, we expect a negativity bias leading to a greater negative effect on birth events because of retrenchment in family allowances with respect to smaller positive effects related to expansive family policies.

Our analysis relies on a dataset merging microdata from the European Union Statistics on Income and Living Conditions (EU-SILC) survey for women aged 18-45 years old interviewed between 2004 and 2020 and aggregate statistics on generosity and universalism of public spending on family-related cash transfers collected for each year and country using the whole EU-SILC samples. For each year and country, relevant changes in the family/children-related allowances are identified as (at least) 10% increases or decreases in the average cash benefit or the share of recipient households with respect to the year before. The event of any relevant change in the provision of family/children-related allowances from two years before the birth to one year before the birth is attributed to the sampled women at the year of interview, thus when the presence of newborns is recorded. This time lag is crucial as it enables us to observe the effect of aggregate changes in social expenditure on birth events in the short-term avoiding any simultaneity bias. Thus, the described data-driven approach allows us to analyze (almost) all European Union countries over a long period of time which includes policy and fertility changes.

Our research contribution is twofold. First, we expand the extensive literature on the effects of welfare policies on fertility using an innovative data-driven approach to identifying relevant changes (seemingly policy reforms) in family/children allowances at national level. Based on a combination of aggregate statistics and microdata from the EU-SILC databases, this approach allows us to focus on a wide range of years and countries instead of almost exclusively considering one new or existing policy in one national welfare system. Secondly, the existing literature predominantly examines expansive reforms in family support, with few exceptions (e.g., González and Trommlerová, 2023). Notably, the social investment agenda has argued that expanding family policies and services positively affects female labor force participation, which, in turn, should also increase fertility (Jenson and Mahon, 2022). While there is strong evidence for the positive relationship between social investment policies and female employment, the relationship with fertility is less clear, especially if family policies undergo austerity measures or are discontinued. To the best of our knowledge, this study is the first to examine the effect of retrenchment changes on fertility across a wide range of years and countries, aiming to assess the presence of an asymmetric reaction in terms of birth events. Indeed, as suggested by the behavioral literature in general and highlighted by González and Trommlerová (2023) in a specific family allowance reform, the magnitude of positive and negative shocks of welfare spending on fertility may not be the same.

Our findings reveal that, as expected, an increase in the generosity of cash benefits is positively associated with a higher probability of childbirth, while the opposite occurs in case of reducing the universalism of family transfers. Decreasing the benefits generosity or increasing their universalism also influence fertility in line with expectations (negative and positive, respectively), but they are statistically insignificant. As a consequence, our main analysis highlights that, on the one hand, individuals react asymmetrically to opposite changes in the provision of family allowances. This means expanding family allowances increases fertility and retrenchment has the opposite effect. On the one other hand, the significance of this asymmetry depends on the type of expansion and retrenchment. While increasing generosity of family transfers has significant positive effects, retrenchment of universalism has significant negative effects. Yet, the opposite is not significant. Results of the heterogeneity analysis however reveal that the asymmetric reaction to changes in the provision of family allowances is not the same across the population but differs according to socio-economic characteristics of mothers (i.e. number of children, household income level, employment status).

The rest of the paper is organized as follows. Section 2 provides a review of the literature on the effect of family policies on fertility. Section 3 describes the datasets used and provides some descriptive statistics. Sections 4 and 5 present respectively the econometric method adopted and the results. The last section concludes and discusses policy implications arising from the analysis.

2. Literature review

Fertility behavior is at the heart of many issues central to the viability of a society. Consequently, it is not surprising that the study of fertility has become a focal point in various social sciences beyond demographics, such as economics and social policy. Central to all these studies is the basic assumption that policies matter, and in particular family policies. Although this hypothesis is not new, it cumulated in the social investment theory that more generous and universal family support has a positive effect on fertility (Jenson and Manson 2022). While more normative studies discuss the limitations and risks of pro-natal policies (Van de Kaa 2006), we concentrate on empirical analyses that study the relationship of family policies and fertility. Findings from these works vary according to the type of policy being considered, the country or set of countries being analyzed, or whether one is investigating complete fertility rather than the timing of births. In this review we will focus only on quantitative multivariate studies that investigate the relationship of family policies and fertility to ensure some basic comparability of assumptions and findings.

Based on the differences by policy type, a distinction can be made between family-friendly labor market policies (such as parental leave and childcare) and family/children allowances to address income inequality and child poverty (such as cash transfers, tax benefits). Work-related policies studies have reached different conclusions, with some finding positive but modest effect on fertility, and others failing to find any substantial evidence. Hoem (1993) reports a positive and significant effect of the parental leave reform approved in Sweden in 1980. A study conducted on parental leave in Sweden and Finland comes to similar conclusions (Ronsen, 2004). No effect for maternity leave is instead found in a study based on Canadian data (Zhang et al., 1994). While Hoem (2001) points out the absence of overall effects of the changes on fertility following the changes in Austria of parental leave in the mid-1990s, Lalive and Zweimuller (2009) note instead that the extension of parental leave in 1990 in Austria increased the probability of having another child, both in terms of time to birth and completed fertility but that the 1996 reduction had no such effect. Kalwij (2010) highlights that a 10% increase in maternity and parental leave benefits results in about a 3.2% reduction in childlessness at ages 36–40 but has no significant effect on completed fertility. Ferrarini (2006) finds that the total fertility rate is positively associated with generous family policies. Differentiated by general family support which tends to promote gendered division of labor, and dual earner support such as shared parental leave entitlements, both policy types have a similar effect on TFR. Hence, family policies increase TFR, but the type of support will have different labor market outcomes. In an updated study (Wesolowski and Ferrarini 2018), covering the years 1995-2010, the positive fertility effects only remain for earner-carer support policies (maternity and shared parental leave). In a systematic review focused on experimental or quasi-experimental studies, Thomas et al. (2022) find mixed effects of parental, maternity, and paternity leave policies on fertility that include positive, negative, and null impacts. However, they suggest that leave policies appear to increase fertility when changes to benefits are generous, particularly through effects on future children rather than the immediate effect on the current child. In a very recent review of the literature, again on experiments and quasi-experiments in low fertility contexts, Hart et al. (2024) highlight that expansions of parental leave rights had positive effects on fertility. Cost and availability of childcare also produce mixed effects on fertility rates, with positive and significant effects in some studies (Del Boca, 2002; Castles, 2003; DiPrete et al, 2003) and zero effects reported in other cases (Hank and Kreyenfeld, 2003; Andersson et al., 2004; Ronsen, 2004).

Tables 1 collects a set of studies on the effect of family allowances (such as cash transfers, tax benefits, tax relief) on fertility, building on summaries already made by Gauthier and Hatzius (1997) and Gauthier (2007). Earlier findings on the effect of family and child benefits aligned with neoclassical economic theory. They suggested a favorable and significant effect of monetary transfers on fertility. This is reviewed for both single country and cross-national studies. Buttner and Lutz (1990) for instance report a statistically significant positive effect of a pronatalist policy introduced in Germany in 1976 on the birth rate up to 5 years after its implementation. A French study (Laroque and Salanie, 2004) states that cash benefits positively influence the probability of having a first birth, although the effect on third births seems to fade. Other two research studies in Canada (Zhang et al., 1994; Milligan, 2005) report that family allowances have a positive effect on fertility. While two cross-national studies, Blanchet and Ekert-Jaffe (1994) on 11 European countries and Gauthier and Hatzius (1997) on 22 OECD countries, show a positive and significant effect of cash transfers on fertility, although small in the second case.

More recent studies have instead presented contradictory evidence, with some scholars continuing to emphasize a positive association between family allowances and fertility, while others find no significant relationship or even revisit previous contributions to challenge their outcomes. For instance, Crump et al. (2011) assert that the results of Whittington et al. (1990) are not robust to more general measures of child tax benefits. In another example, Francesconi and Van der Klaauw (2007) find a counterintuitive negative effect of tax credits on demographic response. In the U.S., expansions in the earned income tax credit led to an extremely small reductions in higher order fertility among white women and no significant effects for non-white women (Baughman and Dickert-Conlin, 2009). Along the same lines, Riphahn and Wiynck (2017) find that the effects of the 1996 reform of the German child benefit program are not statistically significant on low-income couples. The findings of Garganta et al. (2017) differ between childless families and those with children, suggesting a significant positive effect on fertility in families with at least one child, but no significant effect on childless families. In contrast, Wood, Neels, & Vergauwen (2016), focusing on European countries, highlight how family allowances are crucial determinants in the decision to have a second child, particularly among low educated mothers. Milovanska and Farrington (2016) investigated the effect of family allowances in Switzerland and found that higher child benefits incentivized parents to have more children. Other two studies conducted in Israel (Cohen et al., 2013) and Spain (González, 2013) report a positive response of child allowances on fertility. Finally, a recent study conducted in Spain reveals a time-varying effect associated with the introduction the maternity grant, a one-time payment of €2,500 for every child born or adopted. A three percent increase was observed at the time of the policy's introduction, followed by a four percent increase upon the announcement of its repeal, and finally, a six percent decrease in birth rates upon the policy’s actual termination (González and Trommlerová, 2023).

Table 1: Overview of studies on the effect of family allowances on fertility by publication year

Authors	*Methods*	*Countries*	*Years*	*Data*	*Policy type*	*Effect on fertility*
Buttner & Lutz (1990)	Age-period-cohort analysis	Germany	1964-1987	Macro-level data	Pronatalist policy	Statistically significant positive effect of policy on birth rate up to 5 years after implementation
Whittington et al. (1990)	General least-squares regression	USA	1913-1984	Macro-level data	Tax relief, real tax value of the personal exemption	Personal exemption has a positive and significant effect on the birthrate
Blanchet & Ekert-Jaffe (1994)	Ordinary least squares regression and two stage least squares regression	11 European countries	1969-1983	Macro-level data	Cash transfers	Positive and significant effect policy
Zhang et al. (1994)	Generalized least squares	Canada	1921-1981	Macro-level data	Tax exemption, child tax credit, family allowances.	Significant and positive effects on fertility.
Gauthier & Hatzius (1997)	Pooled cross-national and time-series regression	22 OECD countries	1970-1990	Macro-level data	Family cash benefits	Small positive effect of cash benefits on fertility
Laroque & Salanie (2004)	Log-likelihood function and probit model	France	1999-2000	Micro-level data (LFS)	Cash benefits	Cash benefits influence the probability of having a first birth, but not the probability of having a third birth
Milligan (2005)	Probit regression	Canada	1988-1997	Micro-level data	Allowance for newborn children	The cash benefits increase the probability of having a second child by 20.5 percentage point
Bjorklund (2006)	Difference-indifferences approach	Sweden	1960-1980	Macro-level data	Overall measure of family policy measured	Positive effect of family policy on fertility, stable fertility for women born 1930-60 could be explained by other factors
Ferrarini (2006)	Pooled cross-national and time-series regression	18 OECD countries	1970-2000	Macro-level data	Dual Earner / General Family support as net generosity	Both policy types positive effects on TFR
Francesconi & Van der Klaauw (2007)	Simple extension of a difference-in-difference	UK	1991-2001	Micro-level data	Working Families Tax Credit (WFTC)	Negative fertility responses
Baughman & Dickert-Conlin (2009)	Weighted least-squares (WLS)	USA	1990-1999	Micro-level data	Expansions in the earned income tax credit	Extremely small reductions in higher order fertility among white women. No significant effects for non-white women
Crump et al. (2011)	OLS, Prais-Winsten FGLS, general least-squares regression	USA	1913-2005	Macro-level data	Tax benefit: Tax relief, Real tax value of the personal exemption	Whittington et al. (1990) are not robust to more general measures of child tax benefits. No evidence of child tax benefits effect.
Cohen et al. (2013)	Linear probability model, fixed-effects model	Israel	1999-2005	Micro-level data	Child subsidies	Positive, statistically significant, and economically meaningful price effect on overall fertility
González (2013)	Regression discontinuity-type design	Spain	2000-2009	Micro-level data	Universal child benefit	Positive and significant effect on fertility
Laroque, G., & Salanié, B. (2013).	Discrete-choice model	France	1997-1999	Micro-level data	Additional and unconditional child credit	Financial incentives have had a significant effect on fertility decisions in France, strongest for the third child
Milovanska-Farrington (2016)	Instrumental variable approach	Switzerland	2004-2016	Micro-level data	Family allowances	Higher child benefits incentivize parents to have more children
Wood et al. (2016)	Discrete-time hazard models	7 European countries	1970-2002	Micro-level data	Family allowances, childcare enrollment	Family policies, such as family allowances and childcare provisions, are generally positively related to second birth rates.
Garganta et al. (2017)	Diff-in-diff approach	Argentina	2004-2012	Micro-level data	Monthly cash transfers per child	Significant positive effect on fertility in households with at least one child, but no significant effect on childless households.
Riphahn & Wiynck (2017)	Diff-in-diff approach	Germany	1992-1999	Micro-level data	Child cash benefit	Effects on low-income couples are not statistically significant. Positive fertility effects for higher income couples on a second birth
Wesolowski & Ferrarini (2018)	Pooled cross-national and time-series regression	33 advanced economies	1995-2011	Macro-level data	Earner-carer / Traditional-family support as net replacement rate	Earner-carer policies positive effects on TFR
Malak et al. (2019)	Difference-in-differences model	Canada	1988-1997	Micro-level data	Non-taxable cash transfer offered to all resident families in Quebec for each child born or adopted	Significant response to the policy, especially for third-order births or higher, for which the bonus was more generous
Bonner & Sarkar (2020)	Triple difference-in-differences (DDD) strategy	Australia	2003-2008	Micro-level data	Pro-natal policy that provides a one-time payment to families for each new child	Positive impact on fertility, especially among immigrant women with low levels of human capital
González & Trommlerová (2023)	Regression Discontinuity Design, RD–DiD design	Spain	2000-2017	Macro-level data	Universal child benefit	Introduction: 3 percent increase; announcement of cancellation: 4 percent increase; cancellation: 6 percent decrease in birth rates

Source: Elaborations by the authors.

A key concern in the literature is the effect of benefit generosity and universalism on fertility, which have been operationalized in various ways across studies.

Generosity refers to an increase in benefit values, most commonly expressed as net replacement rate in the welfare state literature (Scruggs, 2013). For example, Gauthier & Hatzius (1997) focus on the impact of increasing family transfer generosity expressed as net replacement rate (benchmarked as male earnings in manufacturing). Baughman and Dickert-Conlin (2009) explore expansions in the Earned Income Tax Credit through increases in total benefit amounts as increased generosity. Others measure the extension of eligibility criteria as increased generosity, for instance, extending child benefit from the third child to the second child (Laroque and Salanie, 2004). Milligan (2005) considers the introduction of a policy as increased generosity. Cohen et al. (2013) analyze how variations in family subsidies, with an increase in generosity to existing beneficiaries, affect fertility. We operationalize generosity, similar to Baughman and Dickert-Conlin (2009), as an increase in total benefits per family.

Universalism refers to unconditional benefits based on needs and its positive effects on reducing poverty and inequality have been confirmed in welfare economics (Korpi and Palme, 1998; Brady and Bostic, 2015). On the opposite, targeting refers to means-testing entitlements or stricter eligibility criteria. In our review studies, universalism is seldom discussed in isolation. For instance, Milligan (2005) considers universalism alongside generosity. Gonzalez (2013) focuses on the analysis of the impact of introducing a universal child benefit in Spain in 2007. These studies measure universalism at the policy level, while our study aims to measure universalism at the outcome level (see Korpi and Palme, 1998; Brady and Bostic, 2015). Yet, contrary to Korpi and Palme (1998) we are not using the Kakwani index but use the ratio of households receiving family benefits (see details in methods).

Despite the conflicting empirical evidence, some overarching patterns emerge. Transfers to families, such as cash benefits or tax credits/exemptions, tend to have a small positive effect on total fertility as well as on additional children. However, childless women are not incentivized by general child allowances to alter their fertility decision. Likewise, increases in the generosity and universalism of transfers seem to have positive effects on fertility, albeit in many cases modest and varied for different family conditions, such as economic well-being or the presence or absence of other children.

Based on the literature review, there is a main research gap that has not been addressed. Driven by pioneering Nordic countries, family policies have been expanding in scope and generosity across the globe, but most notably within the European Union. Even family policy laggards like Ireland have implemented all core family policies within a decade (Köppe, 2023). Consequently, country-level studies tend to focus on the expansionary policy effects on fertility and much less on the effects of austerity measures. Second, while earlier studies on this issue analyzed macro-level data and TFR, we contribute to the growing studies that use micro-level data and actual birth events. We are the first study that utilizes EU-SILC to estimate the relationship between actual family allowances received and birth events.

Therefore, our study makes a relevant contribution to the literature. We measure the actual family transfers that households receive. This allows us to show not only increases in per capita benefits, but we can also trace the relative decline of these transfers with a focus on generosity and universalism.

3. Data

The analysis relies on a pooled cross-sectional dataset merging individual-level observations and aggregate statistics on family allowances spending. Both datasets are derived from the European Union Statistics on Income and Living Conditions (EU-SILC) survey and cover the period from 2004 to 2020. The EU-SILC dataset provides detailed micro-data on income, labour, and demographic and socio-economic characteristics at both individual and household level. For this reason, it is a powerful instrument for a comparative analysis of living conditions and receipt of cash social transfers in Europe. The database covers the 27 EU-Member countries and 4 EU Associate Members (Iceland, Norway, Serbia, Switzerland and UK). In this analysis we focus on Iceland and EU-Member countries, excluding Malta and Romania because of missing full information about age of respondents, so that our sample refers to a set of 26 European Union countries. As for the individual-level database, in line with the bulk of the existing literature, we focus on women aged 18-45 years old. Aggregate statistics on generosity and universalism of family/children-related cash transfers are instead collected for each year and country using the whole EU-SILC samples. Within the EU-SILC survey, these allowances include numerous social transfers in the event of childbirth or children living in the home, such as birth grants, family or child allowances, family leave benefits (maternity, paternity, parental), and other cash benefits. For the sake of brevity, we refer to this category of allowances with the label ‘family allowances’ henceforth. As usual in studies relying on EU-SILC data, individual sample weights are adopted when elaborating all descriptive statistics and estimates.

Table 2 presents, for a selection of years (2005, 2010, 2015 and 2020), the number of women in our sample and the share of those who report a new birth in that year. To be noted, instead of considering the number of children born, we count the number of births so that twin births are counted as one.

This operationalization aims to measure the decision to have an (additional) child, but not the randomness of multiple births. Two interesting aspects are evident from Table 2: First, birth rates are decreasing since 2010. Second, differences in fertility rates among EU countries have declined, indicating a downward convergence driven by an overall decrease observed in the rest of Europe. Furthermore, the average birth rates do not reveal clear country clusters that would match welfare or family regimes (Figure A1).

Table 2: Share of women aged 18-45 who gave birth during a selection of years

		2005			*2010*			*2015*			*2020*	
Country	*Women 18-45*	*Women with new birth*	*Share (%)*	*Women 18-45*	*Women with new birth*	*Share (%)*	*Women 18-45*	*Women with new birth*	*Share (%)*	*Women 18-45*	*Women with new birth*	*Share (%)*
Austria	2,522	182	7.2	2,665	194	7.3	2,260	162	7.2	1,935	141	7.3
Belgium	2,444	261	10.7	2,770	240	8.7	2,525	221	8.8	2,639	227	8.6
Bulgaria	2,151*	194*	9.0*	2,657	133	5,0	1,801	74	4.1	2,005	100	5.0
Cyprus	2,335	175	7.5	2,237	107	4.8	2,395	142	5.9	1,927	109	5.7
Czechia	1,878	142	7.6	3,785	231	6.1	2,967	177	6.0	2,664	177	6.6
Germany	5,911	261	4.4	4,689	264	5.6	3,568	180	5.0	6,875	471	6.9
Denmark	2,821	166	5.9	2,341	129	5.5	1,906	91	4.8	1,909	121	6.3
Estonia	2,295	153	6.7	2,583	195	7.5	2,505	197	7.9	2,385	187	7.8
Greece	2,837	165	5.8	3,101	216	7.0	5,512	286	5.2	3,980	155	3.9
Spain	7,533	429	5.7	7,025	448	6.4	5,518	326	5.9	5,958	448	7.5
Finland	4,911	306	6.2	4,182	280	6.7	3,980	287	7.2	3,279	149	4.5
France	4,653	357	7.7	4,629	416	9.0	4,259	361	8.5	3,800	234	6.2
Croatia	n.a.	n.a.	n.a.	4,415	224	6.8	2,597	154	5.9	2,454	101	4.1
Hungary	3,304	212	6.4	4,639	224	4.8	3,151	180	5.7	1,883	110	5.8
Ireland	2,673	147	5.5	1,987	147	7.4	2,340	163	7.0	1,659	85	5.1
Iceland	1,767	132	7.5	166	1,633	10.2	1,512	121	8.0	1,547*	105*	6.8*
Italy	10,911	823	7.5	8,772	497	5.7	6,959	427	6.1	3,731	173	4.6
Lithuania	2,319	143	6.2	2,093	95	4.5	1,477	60	4.1	1,553	58	3.7
Luxembourg	2,022	217	10.7	2,677	255	9.5	1,654	149	9.0	1,397	116	8.3
Latvia	1,840	140	7.6	2,703	203	7.5	2,186	173	7.9	1,810	147	8.1
Netherlands	4,499	521	11.6	4,136	434	10.5	3,470	306	8.8	3,919	206	5.3
Poland	9,597	650	6.8	6,720	585	8.7	5,749	468	8.1	5,753	501	8.7
Portugal	2,390	141	5.9	2,170	109	5.0	3,734	177	4.7	3,910	179	4.6
Sweden	2,708	288	10.6	3,130	355	11.3	2,263	209	9.2	2,207	157	7.1
Slovenia	5,957	284	4.8	5,948	377	6.3	4,721	255	5.4	4,096	224	5.5
Slovakia	3,344	199	6.0	3,569	168	4.7	3,249	155	4.8	2,215	136	6.1
All countries	*93,471*	*6,494*	*6.9*	*95,789*	*8,159*	*8.5*	*84,258*	*5,501*	*6.5*	*75,943*	*4,712*	*6.2*
SD	*2,477*	*176*	*1.9*	*1,854*	*299*	*1.9*	*1,459*	*102*	*1.6*	*1,474*	*117*	*1.5*
Range	*9,144*	*691*	*7.2*	*8,606*	*1,538*	*6.8*	*5,482*	*408*	*5.1*	*5,478*	*443*	*5.0*

*Note: * For Bulgaria, data refer to 2007 first available survey year; For Iceland, data refer to 2018 last available survey year.*

Source: Elaborations by the authors on EU-SILC data, sample weights applied.

Table 3 provides data on the amount of family allowances – at purchasing power parity and inflation-adjusted – reported as the average of transfers received by all households in the sample and as the average for households with children only, for the same selection of years presented in Table 2. Two additional columns in Table 3 show the annual growth rate for the period 2005-2020 of the average benefit amounts. As already mentioned, family policies have expanded in scope and generosity globally, but most notably within the European Union. The average transfer per household increases from €642 in 2005 to €779 in 2020, and the average transfer for households with children increases by almost €700 between 2005 and 2020. Specifically, this increase is driven by many Eastern European countries (especially Poland, Bulgaria, Estonia, Latvia, and Lithuania), which have experienced significant increases over time, indicating a catch-up towards countries with higher spending (β-convergence, Knill, 2005).

Table 3: Mean values of family allowances in the total sample and per recipient

	2005		*2010*		*2015*		*2020*		*2005-2020 annual growth rate (%) – All*	*2005-2020 annual growth rate (%) – With children*
Country	*All*	*With children*	*All*	*With children*	*All*	*With children*	*All*	*With children*		
Austria	1390	4529	1427	4728	1357	4766	1365	4978	-0.1	0.7
Belgium	982	3290	1050	3485	1161	3768	1087	3661	0.7	0.8
Bulgaria	97*	266*	275	731	446	1544	503	1639	27.9	34.4
Cyprus	647	1413	1015	2689	624	1796	594	1711	-0.5	1.4
Czechia	442	1402	476	1598	430	1439	468	1464	0.4	0.3
Germany	949	3513	1004	4014	1090	4424	1309	5440	2.5	3.7
Denmark	488	1923	555	2216	534	2236	646	2882	2.2	3.3
Estonia	419	1148	726	2332	561	1884	1209	4514	12.6	19.5
Greece	166*	264*	173	315	182	642	305	1102	5.6*	21.2*
Spain	83*	170*	132	259	71	131	67	151	-1.3*	-0.7*
Finland	817	3150	887	3516	892	3837	737	3440	-0.7	0.6
France	687	2258	844	3119	860	3132	819	2918	1.3	1.9
Croatia	n.a.	n.a.	335	1003	364	1157	381	1206	0.9	1.3
Hungary	585	1890	836	2639	779	2966	760	2860	2.0	3.4
Ireland	1837	4211	2437	5524	2102	4603	1902	4415	0.2	0.3
Iceland	775	1811	879	2063	638	1774	537**	1489	-2.0**	-1.2**
Italy	254*	744*	273	787	255	786	218	726	-0.9*	-0.2*
Lithuania	169	420	700	2040	372	1104	710	3211	21.3	44.3
Luxembourg	2169	5977	2487	6937	1845	5823	1451	4709	-2.2	-1.4
Latvia	308*	749*	433	1178	436	1401	777	2352	10.2*	14.3*
Netherlands	421	1570	517	2034	452	1841	506	2184	1.3	2.6
Poland	171	445	235	592	309	777	1417	4279	48.6	57.4
Portugal	265*	639*	267	689	161	446	213	625	-1.3*	-0.1*
Sweden	868	3012	818	3105	787	3065	861	3363	-0.1	0.8
Slovenia	778	2114	923	2747	736	2336	825	2692	0.4	1.8
Slovakia	294	746	447	1305	562	1597	591	1688	6.7	8.4
All countries	*642*	*2124*	*775*	*2467*	*692*	*2426*	*779*	*2803*	*1.4*	*2.1*
SD	*525*	*1524*	*594*	*1652*	*491*	*1501*	*445*	*1466*	*11.4*	*15.2*
Range	*2086*	*5807*	*2355*	*6678*	*2031*	*5692*	*1835*	*5289*	*50.8*	*58.9*

*Note: Statistics are expressed in 2015-euro values and adjusted for differences in purchasing power between countries using PPP indexes provided by Eurostat. * Data refer to the first available survey year (2006 for Spain, 2007 for Greece, Italy, Latvia, Bulgaria and Portugal). ** Data refer to the last available survey year (2018). Source: Elaborations by the authors on EU-SILC data, sample weights applied.*

Table A2 in the appendix provides details on the universalism of family allowances, showing the percentage of households receiving benefits in relation to the total number of households in the sample. On average we observe a decline of universalism across the sample countries from 29.6 to 24.3 percent of households receiving family allowances. This is insofar not surprising as fertility has declined as well as households with children.

4. Empirical strategy

The econometric strategy aims to identify the short-term effects of relevant changes in the generosity and universalism of family allowances on fertility. Specifically, we measure the generosity level of these allowances referring to their average benefit amount in the total population. With regards to measuring universalism of these allowances we refer to the share of recipient households in the total population. To account for different economic conditions among countries, different macroeconomic conjunctures between and within countries over the considered period, and different relative importance of cash transfers on household incomes, benefit amounts are also benchmarked against the average household disposable income at national level. Moreover, we differentiate changes in the allowances provision distinguishing between changes in generosity (i.e. benefit amount contractions and increases) and those in universalism (i.e. expansion and reduction of the recipient population with respect to the total one). According to neoclassical theory (Becker, 1991), an increase in transfers to families is expected to raise the fertility rate, as well as an increase in the number of recipients should result in more births. Conversely, a reduction in either benefit amount of family transfers and the recipient population are expected to discourage childbearing.

As for the definition of ‘relevant changes’ in the provision of family allowances, for each year and country, this kind of variations are identified as (at least) 10% increases or decreases in the average cash benefit (*generosity*) and the share of recipient households (*universalism*) with respect to the year before.[13] Table 4 provides an overview of the number of relevant (positive or negative) changes in benefit amounts and the share of recipient households collected between 2004 and 2020 for each country analysed. While the ‘stable situation’ clearly represents the most common one among European countries (73% and 87% of year-country combinations present an overall stable benefit amount and share of recipient households, respectively), several countries increased the generosity and/or the universalism of their family allowances throughout the observation period. Although less common, a number of countries also experienced a reduction in the provision of family allowances during the same period though.

[13] As a sensitivity analysis on the methodological choice adopted to define what changes are ‘relevant’, we replicate the main analysis (presented in Table 5) using 15% instead of 10% as threshold on changes in the average cash benefit or the share of recipient households. Alternative robustness checks may stress our methodological choice in the opposite direction, for instance moving the threshold adopted to define what changes are ‘relevant’ from 10% to 5%. We however decided not to explore this possibility because, even in a scenario without any formal change in the provision of family allowances, small changes in generosity and universalism indicators are expected moving from one EU-SILC wave to another as result of variations in the survey sample. Furthermore, prospect theory suggests stronger reactions to more significant utility changes, hence the 10% threshold sends a stronger behavioral signal.

Tables A3 and A4 expand the information on relevant changes providing the same information as in Table 4 but by year and by crossing possible changes in the provision of family allowances in terms of generosity and universalism, respectively. Table A4 highlights that 69% of the year-country combinations exhibit an overall stability of the family allowances provision (i.e. stable benefit and recipients.

Table 4: Number of relevant change events (10% threshold) in the provision of family allowances from 2004 to 2020 by country

Country	*Decreasing benefit*	*Stable benefit*	*Increasing benefit*	*Decreasing recipients*	*Stable recipients*	*Increasing recipients*
Austria	2	13	1	0	16	0
Belgium	0	16	0	0	16	0
Bulgaria	2	5	6	2	10	1
Cyprus	5	8	2	4	9	2
Czechia	2	11	2	2	13	0
Germany	2	12	1	0	15	0
Denmark	1	14	1	0	16	0
Estonia	4	7	5	0	16	0
Greece	2	7	7	1	9	6
Spain	7	6	3	4	9	3
Finland	0	16	0	0	16	0
France	1	14	1	0	15	1
Croatia	1	8	1	1	9	0
Hungary	2	12	1	1	14	0
Ireland	1	14	1	0	16	0
Iceland	4	8	2	2	12	0
Italy	1	14	1	0	16	0
Lithuania	3	6	6	2	8	5
Luxembourg	0	15	1	2	13	1
Latvia	1	10	4	0	15	0
Netherlands	0	14	1	0	15	0
Poland	1	8	6	0	12	3
Portugal	3	11	2	3	13	0
Sweden	2	14	0	0	16	0
Slovenia	1	14	0	2	12	1
Slovakia	1	11	3	1	14	0
All countries	*49*	*288*	*58*	*27*	*345*	*23*

Source: Elaborations by the authors on EU-SILC data, sample weights applied.

The econometric analysis is based on logistic regressions, where the dependent variable is a binary variable that takes the value 1 if a woman has given birth and 0 otherwise. Within the EU-SILC sample, births are recorded at the time of the interview (time t), while the "changes" in the provision of family allowances are assessed by comparing the levels of generosity and universalism of family allowances in the year before (time t-1) with those levels reported at the country level two years prior (time t-2). For example, if we consider births in 2016, these will be studied in relation to changes in the generosity and universalism of family allowances observed in 2015 (which refer to the difference between the average cash benefit in 2015 compared to 2014 and the share of recipient households in 2015 compared to 2014). This time lag is crucial as it allows us to observe the effect of aggregate changes in social spending on birth events in the short term while avoiding any simultaneity bias, considering the time lag of birth decisions, which includes a typically 9-month “waiting period”. We refer to the 'short term' because we study how fertility behaviors are influenced by observed changes in the provision of family allowances in the year before. Although the 'short term' approach may exclude cases of births related to changes in family allowances reported in subsequent years, the likelihood of further external influences in a more medium-long-term scenario significantly increases. Furthermore, as we want to assess the potential asymmetry in fertility reactions due to relevant changes in family allowances, this kind of reaction is expected to be observed in the short term rather than later in the more distant future.

The specification of the basic Logit model here adopted is as follows:

$$B = \alpha + \beta CHANGE + \gamma \boldsymbol{X} + \theta C + \tau T + \varepsilon,$$

where B represents the probability that a woman gives birth in the year of the interview, $CHANGE$ is a set of dummy variables indicating either a positive or negative change (with respect to stable condition) in average benefit amounts or the share of recipient households of family allowances, and X is a vector of demographic, household, and economic characteristics of women. In particular, the vector X contains variables on age, education level, employment status, marital status, citizenship, number of children already within the household, presence of household members with disability, employment status of other household members, the logarithmic transformation of the total household equivalised disposable income,[14] and the household tenure status. The model also includes an intercept term (α) and both country and year fixed effects (C and T respectively) to account for unobserved heterogeneities across space and time, whereas the error term is represented by ε. Table A5 in the Appendix provides a detailed description of the variables used in the analysis, while Table A6 in the Appendix presents sample descriptive statistics for a selection of the years considered.

According to the set of dummy variables composing the variable $CHANGE$ in the model specification defined above, we distinguish two different models: i) changes in the benefit amount (Model 1); ii) changes in the share of recipients (Model 2).

Finally, we explore potential heterogeneities in the estimated results by some women’s demographic and household characteristics. Specifically, we investigate the results heterogeneity by number of children already present in the family, group of total household equivalized disposable income, and

[14] As usual in studies adopting the EU-SILC data, we adopt as income definition the total household equivalised disposable income, where the equivalence scale is the modified-OECD scale. The modified-OECD scale assigns a value of 1 to the first adult, 0.5 to every other member aged 14 or above, and 0.3 all members aged less than 14. The same equivalised income value is then assigned to each member of the household.

by woman's working status.[15] Especially, the latter is highly relevant in the literature (see Finch and Bradshaw, 2021:653). Since family policies that encourage female employment had been associated with higher fertility, we control for female employment status in our models to ensure we can isolate the family policy effect from employment effects. Furthermore, Finch and Bradshaw (2021) highlight the different effects of in-cash allowances and provision of child-related services like childcare. Since our data only contains data on family allowances, we cannot control for transfers nor services. For the sake of brevity, while commented in Section 5.1, estimation results of the heterogeneity analysis are all provided in the Appendix.

5. Results

Table 5 reports estimated marginal effects of a change in the family allowances provision (change in their generosity and universalism separately) on the probability of observing a woman giving birth. In line with the cumulative evidence reported by Hart et al. (2024) and Finch and Bradshaw (2021), an increase in benefit generosity shows a significant and positive effect on the likelihood of childbirth by 0.5%. Despite this coefficient may appear as limited, considering an average birth rate of 7.1% in the total sample, it means that a relevant increase of the family allowances generosity engenders a 7.0% increase of the birth rate on average, thus representing a demographically relevant encouragement to new births. On the contrary, a benefit retrenchment has no significant effect on the short-term probability of having a child. Hence, we find no evidence for negativity bias, when it comes to benefit cuts and fertility.

As for changes in the family allowances universalism, a decrease in the share of beneficiaries within the total population of households negatively influences the probability of giving birth, while an increase in the share of beneficiaries does not seem to have a statistically significant influence.

The evidence presented by Table 5 confirms that individual reactions in terms of fertility to changes in main features of family allowances is therefore asymmetrical. Conversely to expectations, reactions to changes in transfers generosity seem to report a positivity bias, as the positive growth of birth rates associated to a benefit increase is five times higher (and statistically significant) than the negative growth of birth rates due to a benefit retrenchment. More in line with the literature, changes

[15] Further elaborations of the authors on available data, not provided here for the sake of brevity, have been made also looking at heterogeneous effects by women's age, education level and citizenship. More details are available upon request to the authors.

in universalism follow a negative bias instead with an estimated coefficient 2.5 times larger (and statistically significant) than the one related to a universalism increase.

As a robustness check, we are moving the threshold that defines a change from 10% to 15%. Besides providing a check on the basic model, this allows us to broaden our understanding of effects of changes in family allowances on fertility towards a perspective that contemplates more substantial variations in generosity and universalism and how these can influence fertility choices. The results of this sensitivity analysis, presented in Table A11 in the Appendix, generally confirm the direction of the relationships between generosity/universalism of family allowances and fertility observed in our main analysis. Estimated coefficients overall present greater magnitudes and are always statistically significant. While the positivity bias in the generosity of transfers remains unchanged compared to the main analysis (although reducing its magnitude ratio), the asymmetric reaction in universalism changes. With the threshold at 15%, a higher magnitude is observed in increases in the number of recipients compared to an equal reduction in the beneficiary audience. This discrepancy compared to the main analysis may be linked to the fact that people react differently to the size of changes.

Table 5: Marginal effects on the probability of a birth event

	Model 1	Model 2
Decreasing benefit	-0.001	
	(0.001)	
Increasing benefit	0.005***	
	(0.001)	
Decreasing recipients		-0.005***
		(0.002)
Increasing recipients		0.002
		(0.002)
Time FE	Yes	Yes
Country FE	Yes	Yes
Observations	1,382,515	1,382,515
Pseudo R-squared	0.106	0.106

*Notes: Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Estimated models also contain the list of control variables described in Section 4. Full estimates are provided in Table A7 in the Appendix. Source: Elaborations by the authors on EU-SILC data, sample weights applied.*

5.1 Heterogeneous effects

In fertility research and family dynamics studies, it is crucial to account for heterogeneities in the family composition. Family-oriented policies can indeed have a differential effect on fertility decisions, based on the specific individual or family characteristics (Wood et al., 2016; Riphahn & Wiynck, 2017). One aspect particularly explored in the literature concerns the influence of policies on the desired number of children and, more in general, on the number of children living within a household (Laroque & Salanie, 2004; Garganta et al., 2017; Malak et al., 2019). Policies encouraging fertility may indeed have different effects accordingly to the number of children a household already have.

Table A8 focuses on this potential heterogeneity highlighting that relevant changes in the provision of family allowances affect differently women in the sample according to the number of children they already have before the birth event. Changes in family allowances tend to report coefficients with greater magnitudes among women who already have at least one child. In particular, women already having two or more children within the household appear as the most sensitive to changes in family allowances in general. Childless women, on the other hand, are positively influenced by an increase in benefits and negatively affected by a reduction in recipients. However, they also exhibit results that diverge from the expectations of economic theory: although not significant, an increase in generosity is associated with a positive coefficient sign, while an increase in universalism shows a negative sign. The decision to have the first child, therefore, seems to be influenced by other and broader socio-economic and cultural factors, which marginally consider the response to policy changes. Conversely, women with children show a greater sensitivity to changes in family allowances. This finding, already emphasized in previous studies (Wood et al., 2016; Milovanska and Farrington, 2016; Garganta et al., 2017), could be interpreted as an indication that larger families tend to perceive themselves as more vulnerable and needy of a financial support in order to expand further. Moreover, women with children have a much clearer anticipation of the cost and benefits associated with each additional child. Thus, an increase in generosity is potentially already experienced for the child(ren) in the household at t-1, while childless women operate in an anticipation of how the family allowances will affect the household income without the lived experience.

As for the heterogeneous effects by tertile group of the household equivalised disposable income, Table A9 in the Appendix shows that fertility decisions of women living in low or middle-income households are the only ones affected by changes in the generosity of family allowances. This evidence is in line with expectations, as fertility decision of high-income families may be less sensitive to welfare state support, especially when it concerns means-tested benefits. When it comes to universalism, the perspective reverses, because only women belonging to the highest income tertile respond significantly and negatively to reductions in the share of beneficiaries. This confirms Korpi and Palme (1998) that households across the income distribution benefit from universal benefits. The strong negative effects indicates that higher income earners react negatively to stronger targeting of family allowances (e.g. stronger means-tests), while low-income families are less likely to be affected by a reduction of the universalism of family allowances.

Finally, regarding the heterogeneity analysis based on women's working status, Tables A10 in the Appendix reveals interesting differences in fertility reactions to significant changes in family allowances. Specifically, we distinguish between full-time employment, part-time employment, and non-employment (i.e. students, unemployed, or other inactive status). In line with predictions from economic theory, non-employed women represent the category most affected by changes in both the generosity and universalism of family allowances. Their fertility reactions reflect the direction of changes with a particular emphasis on negative variations, so that a decrease in benefit amounts or the share of recipients leads to a strong and significant reduction in birth events, while the opposite (but not significantly) occurs when benefits or the number of recipients increase. Fertility decisions of women in full-time employment are instead only (and positively) influenced by an increase in the generosity of transfers. For part-time working women, we do not observe significant coefficients. Yet, Finch and Bradshaw have highlighted how social investment policies that emphasize social services such as childcare provision, full day schools or flexible working arrangements are more relevant for working mothers. Hence, our narrower focus on family allowances is not fully capturing welfare policy interventions with an effect on fertility decisions.

6. Conclusions and policy implications

In the context of growing interest in family support policies as a potential tool to counteract declining fertility rates, this study explores the influence of changes in the generosity and universalism of family and child allowances on the likelihood of having a child.

Our results indicate that an increase in the generosity of cash benefit is associated with a significant and positive increase in the probability of childbirth. On the other hand, a decrease in the share of recipients engenders a negative and significant influence to the probability of giving birth. Decreasing the benefits generosity or increasing their universalism also have effects on fertility which appear in line with expectations (negative and positive, respectively), but they are statistically insignificant. Consequently, our main analysis highlights two key findings. First, women react asymmetrically to changes in the provision of family allowances. Second, the asymmetrical reaction is shaped by the type of welfare change, i.e. generosity versus universalism. In particular, the fertility increase due to a generosity increase is about five times greater than the opposite fertility decrease due to a benefit retrenchment, whereas the fertility reaction to a decrease of transfers universalism is about 2.5 times greater than the one due to the opposite event.

Results of the heterogeneity analysis however reveal that the asymmetric reaction to changes in the provision of family allowances is not the same across the population but differs according to characteristics of mothers (i.e. number of children, household income level, employment status). This evidence highlights that a careful analysis of the needs and preferences of different population segments can aid in designing more effective interventions, thereby enhancing the overall impact of policies on fertility.

Our analysis has a few limitations, due to the nature of the EU-SILC data. First, we have measured changes in generosity and universalism at the aggregate household level. This means we could not measure policy changes directly. Further research could explore how to measure direct family allowance changes at the country level (e.g. OECD family database) have lagged effects on fertility decisions. Second, our measure of universalism is a rather crude measure that is also sensitive to changes in overall fertility. Since we measure universalism across all households, a declining fertility is most likely auto-correlated with a decline in universalism. Further analysis should test for other possible universalism measures like the Kakwani index used by Korpi and Palme (1998). Third, with our longer observation period, we could not use the more fine-grained differentiation between means-tested and contribution-based family allowances, which has been implemented in EU-SILC from 2014 onwards. Hence, a focus on means-tested allowances would provide a more direct measure of universalism. Fourth, we cannot differentiate between policies that encourage overall fertility (e.g. leave policies) and a higher number of children, like promoted by the Hungarian Orban government (see below). At the same time, our focus on allowances omitted social services that were identified to be highly relevant for the fertility decisions of well-educated and working women (Finch and Bradshaw 2021).

We find evidence that the observed asymmetric reaction might be due to potential cognitive biases and/or imperfect information about the provision of social transfers. In our comparison of childless women and women with children, we can show that women with children react more sensitively to changes in generosity, since they have more complete information about the actual social transfers received per child. To be noted, the latter issue can also be exacerbated by a possible incomplete communication by policymakers who may be more interested in emphasizing spending increases and concealing budget cuts. The lack of clarity or incomplete communication can not only confuse beneficiaries but also alter the effectiveness of such policies. Therefore, to maximize the efficacy of family support policies in addressing declining fertility rates, policymakers should prioritize transparent communication and ensure that policy adjustments are carefully crafted to meet the specific needs of diverse groups within society.

Furthermore, several countries in the sample have reduced the generosity of family allowances for children of a higher birth order in our observation period. The United Kingdom (Chzen and Bradshaw 2024) and Italy (Aprea et al. 2024) have introduced punitive social assistance schemes that cap family allowances at two and three children, respectively. Also Germany has reduced the value of their child benefit for the third and fourth child between 1996 and 2022 (Köppe et al. 2024). Hence, our findings point to less effects on fertility regarding large families, but confirm the negative effects on child poverty (Köppe et al. 2024, Chzen and Bradshaw 2024). Yet, our data does not allow to interrogate large families separately, but points to a focus on child poverty alongside fertility.

In conclusion, following our results, policymakers should carefully weigh the potential outcomes of their choices. So far, the mantra of the social investment agenda had claimed that introducing generous family policies will increase fertility (Finch and Bradshaw, 2021; Jenson and Mahon, 2022). And indeed, increased benefit generosity shows a disproportionately larger impact on fertility than equivalent reductions. Conversely, this also means punitive approaches are less effective in achieving lower fertility with subsequent adverse effects on child poverty (see Chzchen and Bradshaw 2024). Furthermore, prioritizing enhancements to the generosity of transfers over broadening the recipient base could yield more substantial effects on fertility rates. Conversely, during periods of resource constraint, reductions in the universalism of family allowances may prompt more pronounced reactions, suggesting a need for cautious consideration in policy adjustments. In other words, introducing strong means-tests in a periods of austerity, results in significant reductions in fertility. Optimizing the allocation of financial resources towards augmenting the generosity of transfers presents a promising strategy to bolster fertility rates without compromising the inclusivity of support programs. This approach acknowledges the differential impacts of policy changes and underscores the importance of strategic resource allocation in maximizing the effectiveness of family support policies.

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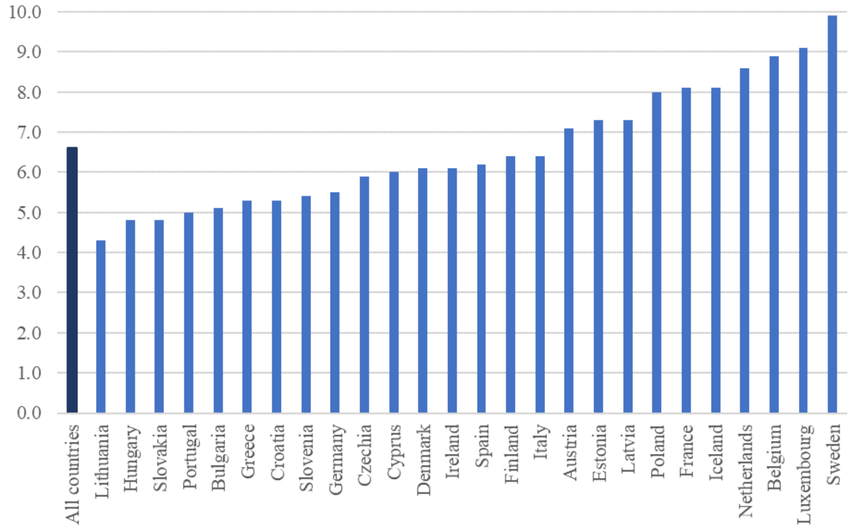
Appendix

Figure A1 – Average birth rate by country for women 18-45 years



Source: Elaborations by the authors on EU-SILC data, sample weights applied.

Table A1 – Average values of family allowances for households with and without children

Country	No children	One child	Two children
Austria	394	4,067	6,714
Belgium	256	2,509	5,348
Bulgaria	28	342	614
Cyprus	154	1,180	2,449
Czechia	36	840	1,106
Germany	305	3,477	5,924
Denmark	55	2,291	4,114
Estonia	87	1,370	2,428
Greece	70	276	622
Spain	51	179	134
Finland	119	3,198	5,496
France	78	1,670	4,515
Croatia	36	449	1,012
Hungary	57	1,103	2,284
Ireland	633	4,624	7,567
Iceland	207	2,051	2,778
Italy	86	544	1,141
Lithuania	62	741	1,036
Luxembourg	595	5,542	10,000
Latvia	81	727	1,239
Netherlands	23	1,414	2,717
Poland	52	432	1,156
Portugal	48	351	693
Sweden	131	3,123	4,988
Slovenia	139	1,585	2,677
Slovakia	70	760	1,103
All countries	152	1,763	3,141
SD	164	1,474	2,550
Range	610	5,363	9,866

Notes: For Spain, data refer to 2006 first available survey year; for Greece, Italy, Latvia, Bulgaria and Portugal to 2007; for Iceland data refer to 2018 last available survey year. Source: Elaborations by the authors on EU-SILC data, sample weights applied.

Table A2 - Share of households receiving family allowances

Country	*2005*	*2010*	*2015*	*2020*	*2005-2020 annual growth (%)*
Austria	32.0	30.75	30.31	28.24	-0.7
Belgium	31.4	32.02	31.71	30.92	-0.1
Bulgaria	20.2	21.68	19.85	20.83	0.2
Cyprus	53.9	46.48	24.68	26.86	-3.1
Czechia	29.3	16.68	12.24	10.95	-3.9
Germany	30.4	28.39	26.70	26.83	-0.7
Denmark	26.1	25.55	24.78	22.31	-0.9
Estonia	34.0	30.27	28.26	26.03	-1.5
Greece	11.7*	11.14	12.72	22.29	5.7
Spain	3.3*	4.64	2.78	2.66	-1.2
Finland	24.8	23.83	22.24	20.48	-1.1
France	25.7	24.83	25.29	22.51	-0.8
Croatia	n.a.	15.79	15.34	12.35	-1.4
Hungary	31.8	34.31	27.40	27.34	-0.9
Ireland	46.1	45.97	44.93	39.89	-0.8
Iceland	30.9	34.08	26.80	19.7**	-2.3
Italy	27.1*	25.54	23.36	20.84	-1.4
Lithuania	15.2	29.88	10.92	27.97	5.3
Luxembourg	37.5	37.52	31.46	30.52	-1.2
Latvia	36.0*	32.24	27.84	30.11	-1.0
Netherlands	27.5	26.75	26.10	23.13	-1.0
Poland	21.0	15.05	11.93	33.36	3.7
Portugal	35.6*	30.36	15.41	14.96	-3.6
Sweden	26.1	23.19	22.93	23.51	-0.6
Slovenia	37.7	34.50	22.64	26.40	-1.9
Slovakia	41.8	39.25	37.51	36.87	-0.7
All countries	*29.6*	*27.7*	*23.3*	*24.3*	-1.1
SD	10.62	9.87	9.02	8.02	2.3
Range	50.60	41.84	42.15	37.23	9.6

*Notes: *For Spain, data refer to 2006 first available survey year; for Greece, Italy, Latvia, Bulgaria and Portugal to 2007; **for Iceland data refer to 2018 last available survey year. Source: Elaborations by the authors on EU-SILC data, sample weights applied.*

Table A3 - Number of relevant change events in the provision of family allowances from 2004 to 2020 by year

Year	*Decreasing benefit*	*Stable benefit*	*Increasing benefit*	*Decreasing recipients*	*Stable recipients*	*Increasing recipients*
2005	2	10	2	1	12	1
2006	3	18	3	0	23	1
2007	4	17	3	1	20	3
2008	4	18	3	2	22	1
2009	2	14	9	1	22	2
2010	1	18	6	0	23	2
2011	3	21	2	4	22	0
2012	7	18	1	7	19	0
2013	5	18	3	1	24	1
2014	4	18	4	3	22	1
2015	4	19	3	2	24	0
2016	2	19	5	1	23	2
2017	2	18	6	1	22	3
2018	1	22	3	2	23	1
2019	2	21	2	1	21	3
2020	3	19	3	0	23	2
All years	*49*	*288*	*58*	*27*	*345*	*23*

Source: Elaborations by the authors on EU-SILC data, sample weights applied.

Table A4 – Share of relevant change events in the provision of family allowances

Change	**Total sample**
Decreasing benefit	12,21%
Stable benefit	73,29%
Increasing benefit	14,50%
Decreasing recipients	6,57%
Stable spending	86,65%
Increasing recipients	6,78%
Decreasing benefit and decreasing recipients change	3.80%
Decreasing benefit and stable recipients change	7.91%
Decreasing benefit and increasing recipients change	0.50%
Stable benefit and decreasing recipients	2.49%
Stable benefit and recipients	68.63%
Stable benefit and increasing recipients	2.17%
Increasing benefit and decreasing recipients change	0.28%
Increasing benefit and stable recipients change	10.10%
Increasing benefit and increasing recipients change	4.12%

Source: Elaborations by the authors on EU-SILC data, sample weights applied.

Table A5 - List of variables used

Variables	*Description*
Dependent variable	
New Birth	Binary variable that takes value 1 for women who give birth in the year in which the EU-SILC interview is conducted and 0 otherwise. Births in the sample are taken at the time of the interview (time t).
Key explanatory variables	
Decreasing benefit Increasing benefit	Binary variables indicating a change in cash transfers on family allowances. Transfers are benchmarked against the national average total disposable household income. If the variation exceeds 10% of the previous year's total transfers, it is considered as a "change". Increasing if the variation is positive and decreasing if negative. The reference category is no "change" i.e., a variation of less than 10% comparing to the previous year's benefit. Family allowances in the sample are taken at the year before the interview (time t-1).
Decreasing recipients Increasing recipients	Binary variables indicating a change in the number of recipients of family allowances. If the variation in recipients exceeds 10% of the total number of recipients in the previous year, it is considered a "change". The reference category is no "change," i.e., a variation of less than 10% from the number of previous year's recipients. Family allowances recipients in the sample are taken at the year before the interview (time t-1).
Spending contraction Spending expansion	Binary variables indicating both a change in cash transfers on family allowances and a change in the number of recipients of family allowances. Transfers are benchmarked against the national average total disposable household income. The reference category is no "change," i.e., a variation of less than 10% comparing to the previous year's benefit and less than 10% from the number of previous year's recipients. Family allowances in the sample are taken at the year before the interview (time t-1).
Control variables	
Age	Discrete variable on women's age. Range between 18 and 45 years.
Education	Binary variable that takes the value 1 for women with upper secondary education or more, 0 otherwise.
Self-Employed Unemployed Student Other status	Binary variables representing the women's employment status. The reference category is Employed
Married	Binary variable taking value 1 for married women, and 0 otherwise.
Migrant	Binary variable taking value 1 for migrant women, and 0 otherwise.
Already one child Already two children Already three or more children	Binary variables representing the number of children of the women sampled. The reference category consists of No children.
Presence of members with disability	Binary variable taking value 1 for women with at least one disabled person in the household, 0 otherwise.
People employed in the household	Discrete variable on the number of other household members employed.
Equivalised income	Continuous variable on the logarithm of household equivalised disposable income.
Mortgage	Binary variable that takes the value 1 for women who have a mortgage, 0 otherwise.

Table A6 - Sample descriptive statistics in a selection of years

Variables	*2007*	*2010*	*2015*	*2020*
Observations	94090	92843	84257	75937
Age (in years)	32.26	32.26	32.50	32.72
Education				
Lower secondary education	19.79%	19.07%	15.86%	13.86%
Upper secondary education	80.21%	80.93%	84.14%	86.14%
Employment status				
Employed	66.49%	67.68%	66.41%	71.40%
Self-employed	5.77%	5.50%	5.82%	5.60%
Unemployed	5.08%	5.37%	7.88%	5.11%
Student	8.58%	9.00%	8.69%	7.50%
Other	14.08%	12.44%	11.21%	10.40%
Marital status				
Not married	50.78%	53.22%	56.08%	58.27%
Married	49.22%	46.78%	43.92%	41.73%
Citizenship				
Local	86.74%	86.08%	91.96%	90.84%
Migrant	13.26%	13.92%	8.04%	9.16%
Number of children				
No children	37.81%	39.12%	39.86%	41.25%
One child	28.22%	28.07%	27.46%	26.12%
Two children	24.66%	23.87%	24.00%	24.15%
Three or more children	9.31%	8.94%	8.68%	8.48%
Presence of members with disability				
No people with disability in the household	94.04%	95.21%	95.53%	95.94%
People with disability in the household	4.96%	4.79%	4.47%	4.06%
Household equivalised income (in €)	14,738.95 €	15,640.85 €	16,052.29 €	19,489.54 €
People employed in the household	1.63	1.65	1.59	1.65
Mortgage				
Without Mortgage	72.15%	68.33%	68.17%	67.06%
With Mortgage	27.85%	31.67%	31.83%	32.94%

Notes: Given that for several countries the earliest available data refer to years after 2005 (For Spain, data refer to 2006 as the first available survey year; for Greece, Italy, Latvia, Bulgaria, and Portugal to 2007), to have comparable statistics, the earliest data in the table do not refer to 2005 but to 2007. Source: Elaborations by the authors on EU-SILC data, sample weights applied.

Table A7 - Marginal effects on the probability of a birth event: full estimates

Variables	*Whole sample* *Spending* change	*Whole sample* *Recipients* change
Age	-0.005***	-0.005***
	(0.000)	(0.000)
Education	0.011***	0.011***
	(0.001)	(0.001)
Employment status		
Self-Employed	-0.002	-0.002
	(0.002)	(0.002)
Unemployed	0.013***	0.013***
	(0.002)	(0.002)
Student	-0.058***	-0.058***
	(0.001)	(0.001)
Other status	0.032***	0.032***
	(0.001)	(0.001)
Married	0.095***	0.095***
	(0.001)	(0.001)
Migrant	0.008***	0.008***
	(0.001)	(0.001)
Number of children		
Already one child	0.010***	0.010***
	(0.001)	(0.001)
Already two children	-0.053***	-0.053***
	(0.001)	(0.001)
Already three or more children	-0.049***	-0.049***
	(0.001)	(0.001)
Presence of members with disability	-0.016***	-0.016***
	(0.002)	(0.002)
People employed in the household	-0.005***	-0.005***
	(0.001)	(0.001)
Equivalised income	-0.003***	-0.003***
	(0.000)	(0.000)
Mortgage	0.024***	0.024***
	(0.001)	(0.001)
Change in family allowances		
Decreasing benefit	-0.001	
	(0.001)	
Increasing benefit	0.005***	
	(0.001)	
Decreasing recipients		-0.005***
		(0.002)
Increasing recipients		0.002
		(0.002)
Time fixed effects	Yes	Yes
Country fixed effect	Yes	Yes
Observations	1,382,515	1,382,515
Pseudo R-squared	0.106	0.106

*Notes: Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Source: Elaborations by the authors on EU-SILC data, sample weights applied.*

Table A8 - Marginal effects on the probability of a birth event by number of children already within the household

	Whole sample		No children		One child		Two or more children	
	Model 1	Model 2	Model 1	Model 2	Model 1	Model 2	Model 1	Model 2
Decreasing benefit	-0.001		0.001		-0.003		-0.003*	
	(0.001)		(0.002)		(0.003)		(0.002)	
Increasing benefit	0.005***		0.003**		0.006**		0.005***	
	(0.001)		(0.002)		(0.003)		(0.002)	
Decreasing recipients		-0.005***		-0.005*		-0.004		-0.004
		(0.002)		(0.003)		(0.003)		(0.003)
Increasing recipients		0.002		-0.002		0.006		0.007***
		(0.002)		(0.002)		(0.004)		(0.003)
Time FE	Yes		Yes		Yes		Yes	
Country FE	Yes		Yes		Yes		Yes	
Observations	1,382,515		584,787		379,775		417,953	
Pseudo R-squared	0.106	0.106	0.143	0.143	0.092	0.092	0.065	0.065

*Notes: Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Estimated models also contain the list of control variables described in Section 4. Source: Elaborations by the authors on EU-SILC data, sample weights applied.*

Table A9 - Marginal effects on the probability of a birth event by household income group

	Whole sample		Low income		Middle income		High income	
	Model 1	Model 2	Model 1	Model 2	Model 1	Model 2	Model 1	Model 2
Decreasing benefit	-0.001		-0.004*		-0.002		0.003	
	(0.001)		(0.002)		(0.002)		(0.002)	
Increasing benefit	0.005***		0.004**		0.004*		0.003	
	(0.001)		(0.002)		(0.002)		(0.003)	
Decreasing recipients		-0.005***		-0.003		-0.003		-0.011***
		(0.002)		(0.003)		(0.003)		(0.004)
Increasing recipients		0.002		0.003		-0.001		-0.005
		(0.002)		(0.002)		(0.003)		(0.004)
Time FE	Yes		Yes		Yes		Yes	
Country FE	Yes		Yes		Yes		Yes	
Observations	1,382,515		460,723		461,182		460,610	
Pseudo R-squared	0.106	0.106	0.109	0.109	0.114	0.114	0.121	0.121

*Notes: Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Estimated models also contain the list of control variables described in Section 4. The group 'Low income' considers women living in households reporting an income lower than 8,519 €. The group 'Middle income' between 8,520 € and 18,573 €. The group 'High income' higher than 18,573 €. Source: Elaborations by the authors on EU-SILC data, sample weights applied.*

Table A10 - Marginal effects on the probability of a birth event by woman's working status

	Whole sample		Full-time worker		Part-time worker		Non-employed	
	Model 1	Model 2	Model 1	Model 2	Model 1	Model 2	Model 1	Model 2
Decreasing benefit	-0.001		0.001		-0.005		-0.006***	
	(0.001)		(0.002)		(0.003)		(0.002)	
Increasing benefit	0.005***		0.003**		0.005		0.002	
	(0.001)		(0.001)		(0.003)		(0.002)	
Decreasing recipients		-0.005***		-0.001		-0.008		-0.007**
		(0.002)		(0.002)		(0.006)		(0.003)
Increasing recipients		0.002		-0.000		-0.001		0.003
		(0.002)		(0.002)		(0.005)		(0.003)
Time FE	Yes		Yes		Yes		Yes	
Country FE	Yes		Yes		Yes		Yes	
Observations	1,382,515		658,950		192,289		531,276	
Pseudo R-squared	0.106	0.106	0.131	0.131	0.145	0.145	0.117	0.117

*Notes: Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Estimated models also contain the list of control variables described in Section 4. Source: Elaborations by the authors on EU-SILC data, sample weights applied.*

Table A11 - Marginal effects on the probability of a birth event: alternative definition of relevant change of the family allowances provision

	Model 1	Model 2
Decreasing benefit	-0.004**	
	(0.002)	
Increasing benefit	0.007***	
	(0.001)	
Decreasing recipients		-0.004**
		(0.002)
Increasing recipients		0.006***
		(0.002)
Time FE	Yes	Yes
Country FE	Yes	Yes
Observations	1,382,515	1,382,515
Pseudo R-squared	*0.106*	*0.106*

*Notes: Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Estimated models also contain the list of control variables described in Section 4. Source: Elaborations by the authors on EU-SILC data, sample weights applied.*