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Who is admitted for bipolar disorder, and where: a 25-year national analysis of hospital discharges in Chile

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Abstract

Purpose.

Admission for a severe mental disorder reflects not only clinical need but the alternatives a system offers outside hospital. We examined how the volume, age and sex composition and territorial distribution of bipolar disorder (BD) admissions changed in Chile over 25 years, and whether the 2013 incorporation of BD into the Explicit Health Guarantees scheme altered them.

Methods.

Time-series study of all public and private hospital discharges in Chile, 2001–2025, with a primary diagnosis of ICD-10 F31 (59,208 episodes). Rates from 2002 used official population projections and were age-standardised (WHO world standard). Trends were modelled by joinpoint regression, sex-by-age patterns by quasi-Poisson regression, territorial inequality by the slope and relative indices of inequality, and the guarantee by segmented regression under alternative pre-intervention specifications.

Results.

The age-standardised rate rose from 10.0 to 16.1 per 100,000 (+ 1.9% per year; 95% CI 1.4–2.5), accelerating to 6.6% after 2021. Rates roughly tripled at ages 10–19 and more than doubled at 20–29 while falling after age 70. Women were admitted at about twice the rate of men (1.98; 1.92–2.04), but that gap widened only after age 60. A fivefold gap between extreme regions persisted while population-weighted inequality narrowed. The step reduction attributed to the guarantee under a single pre-intervention slope (RR 0.83; 0.73–0.95) attenuated to 0.94 (0.88–1.01) once the decline under way since 2010 was modelled.

Conclusion.

Inpatient care for BD shifted toward adolescents and young adults and remains territorially concentrated. The apparent effect of the 2013 guarantee is not robust to how the pre-existing trend is specified.

Introduction

Bipolar disorder (BD) involves alternating episodes of mania or hypomania and depression, separated by intervals that may be asymptomatic or subsyndromal [1, 2]. Onset clusters in the second decade of life and the course is typically recurrent, with lasting functional impairment in a substantial proportion of patients [1, 3]; global estimates place incidence as rising among adolescents and young adults, most steeply between ages 20 and 24 [4].

Inpatient admission captures the severe end of this course: episodes in which outpatient management proved insufficient, risk became unmanageable, or no community alternative was available. That makes

discharge records informative about the social organisation of care, and a poor proxy for the occurrence of disorder. A rise in admissions is compatible with more illness, with earlier detection, with more beds, or with lower thresholds for admitting, and administrative data alone cannot separate these. We therefore treat the outcome throughout as inpatient care utilisation. Read that way, national discharge series have documented the redistribution of psychiatric care in several settings: a 28% fall in admissions across NHS hospitals in England between 1998/99 and 2019/20 [5], admissions that were stable or rising in Czechia alongside sharply shorter stays [6], and the reshaping of schizophrenia admissions during deinstitutionalisation in Spain [7]. What such series describe is less the epidemiology of a diagnosis than the capacity of a system to manage crises somewhere other than a hospital bed.

Chile is an informative setting for that question on three counts. First, its health system is segmented: a public insurer (FONASA) covers most of the population and private insurers (ISAPRE) cover a smaller, higher-income minority, so the inpatient population is socially selected before any clinical threshold applies. Second, its territory is unusually heterogeneous, with regions ranging from about one hundred thousand to more than eight million inhabitants, and, as across Latin America, psychiatric beds and community teams are distributed very unevenly, with acute capacity for children and adolescents particularly scarce [8]; regional and gender disparities in depression and in healthy life expectancy have recently been described for the same country [9]. Third, in July 2013 treatment of BD in people aged 15 and over was incorporated into the Explicit Health Guarantees scheme, which establishes legally enforceable guarantees of access, timeliness, financial protection and quality [10]. Chilean mental health policy has been evaluated with segmented regression before, for the national mental health plans of 1993 and 2000 [11], and the approach has been applied to service use elsewhere in the region [12], which makes it a natural starting point here.

The regional picture is also shifting: Brazil recorded a sustained fall in psychiatric hospitalisations between 2008 and 2022 as community services expanded [13], and Ecuador has described eleven years of discharge data for BD specifically [14]. Long single-diagnosis series nonetheless remain rare in Latin America, and the closest counterpart outside it is a Spanish national analysis of admissions among 10–19-year-olds [15]. National longitudinal evidence for Chile derives mostly from aggregated administrative reports, which cannot support age-standardised comparison across regions or the identification of turning points.

We analysed all national hospital discharges recorded over 25 years to estimate annual crude and age-standardised rates; to describe discharges by sex, age, region, diagnostic subtype, length of stay and insurance; to identify change-points nationally and by sex and age; to quantify territorial disparity using the whole regional distribution; and, as a secondary question, to test whether the 2013 guarantee was associated with a change in the level or slope of rates.

Methods

Design and data source

We conducted a nationwide, population-based study with repeated annual cross-sections and time-series analysis, using the administrative hospital discharge records compiled by the Department of Health Statistics and Information (DEIS) of the Chilean Ministry of Health. These open, anonymised databases cover all discharges from public and private facilities and code diagnoses according to ICD-10 [16]. The 25 annual files were downloaded on 9 August 2026, when the repository held complete files through 2025 and none for 2026; file sizes, modification dates and SHA-256 checksums are deposited with the analysis code.

The unit of analysis is the discharge episode, not the individual. DEIS releases no patient identifier, so repeat admissions by the same person are counted separately and cannot be linked. Reporting follows the RECORD extension of STROBE for routinely collected health data [17].

Case definition, variables and harmonisation

All discharges recorded between 2001 and 2025 with a primary diagnosis of F31.x were included (Fig. S2). Episodes coded F30.x (manic episode) were analysed separately, since an isolated manic episode does not establish a diagnosis of BD. Variables were year of discharge, sex, age group, region of residence, ICD-10 subtype, length of stay, discharge condition and health insurance.

Two features of the source required harmonisation, and both differ from what the available literature on these data would suggest. Age is released only in grouped form, and the grouping is not stable: most years use ten-year bands, whereas 2012 and 2021 use five-year bands closing at "85 and over", so we adopted the finest scheme common to all years, ten-year bands with an open category at 80 and over. One consequence matters for the policy analysis: because 15 is not an available boundary, the population covered by the guarantee cannot be isolated, and the natural internal control of comparing those aged 10–14 with those aged 15–19 is unavailable. Region of residence required no retrospective regrouping; although Los Ríos and Arica y Parinacota were created in 2007 and Ñuble in 2018, DEIS codes region under the current sixteen-region scheme throughout the series, and records for all three appear from 2001 onward. Sex was recorded as text in most years and numerically in 2012 and 2021, and was recoded to a single scheme.

Denominators and standardisation

Population denominators came from the official INE estimates and projections by year, region, sex and single year of age (2017 base) [18]. These projections begin in 2002. The year 2001 is therefore reported for absolute counts and cohort description but excluded from every rate, trend and model; no rate for 2001 can be computed from the official denominators, and we did not substitute an unofficial series.

Annual rates per 100,000 were computed as crude, sex- and age-specific, regional, and age-standardised by the direct method using the WHO world standard population [19] collapsed to the study age bands.

Confidence intervals for standardised rates use the Fay–Feuer gamma method [20], which behaves appropriately when some strata contribute few events; crude rates use Byar's approximation.

Missing and masked records

DEIS masks sex, age or region in a minority of records. Across the F31 series 96.2% of episodes had all three recorded, but completeness varied sharply by year and fell to 55.7% in 2002, so a complete-case analysis would halve the estimated rate for that year and manufacture a spurious rise at the start of the series. Assuming masking is non-differential within a calendar year, we scaled each year's denominator by that year's completeness, which is equivalent to redistributing masked episodes in proportion to the observed structure and preserves the annual total exactly. The uncorrected complete-case analysis is reported as a sensitivity analysis, and completeness by variable and year is given in the supplementary material (Fig. S1).

Statistical analysis

Categorical variables are summarised as frequencies and percentages with the denominator stated for each, since insurance status is unknown for a non-trivial share of records; length of stay is summarised as median and interquartile range.

Trends in standardised rates were analysed by log-linear joinpoint regression, an approach used previously for national self-harm series in this journal [21]. We fitted log(rate) as a piecewise linear function of calendar year by weighted least squares, weighting by the number of events, searching exhaustively over knot positions with a minimum of four points per segment and selecting the number of knots (zero to three) by the Bayesian information criterion. Segment-specific annual percent change (APC) and the average annual percent change (AAPC) are reported with 95% confidence intervals, the latter as a length-weighted combination of segment slopes with variance obtained by the delta method. All segments of all series are reported. This implementation differs from the NCI Joinpoint Regression Program, whose default selects knots by permutation tests [22]; the code is deposited.

Associations of rates with sex, age group, period and region were estimated by regression with the log of population as offset. The Pearson dispersion statistic was 2.67, so quasi-Poisson models were used and the dispersion parameter is reported. Calendar time was modelled as period indicators defined by the joinpoint knots rather than as a single linear term, because a linear coefficient averaged across segments with opposite slopes has no substantive interpretation. Residuals of the aggregated annual series were autocorrelated (Durbin–Watson 0.84, $p < 0.001$), so estimates from annual models carry Newey–West standard errors.

To test whether the widening sex gap was concentrated in the age groups where rates were rising, we compared the female-to-male rate ratio within each age group between 2002–2013 and 2014–2025, and

assessed the three-way age by sex by period interaction against a model containing all two-way interactions, using an F test appropriate to the quasi-Poisson fit.

Territorial disparity was summarised in two ways. The ratio between the highest and lowest regional rate is reported for comparability with earlier work, but because it uses only two of sixteen regions and is sensitive to sampling variation in small populations, we also report the slope index of inequality (SII) and the relative index of inequality (RII), computed on the population-weighted ranking of all regions.

The policy analysis used segmented regression [23] with a level change from 2014 and a slope change thereafter, treating 2013 as a transition year. Because the joinpoint analysis identified a decline beginning in 2010, three years before the intervention, a single pre-intervention slope from 2002 misattributes to the policy a deviation already under way. We therefore fitted the model under several specifications: a single pre-intervention slope, a slope break at 2010, and a pre-intervention window restricted to 2010–2013. The primary analysis was restricted to those aged 20 and over, who are unambiguously covered by the guarantee, with those aged 10 and over reported for comparability with the specification used previously.

Sensitivity analyses combined F30 with F31, excluded the pandemic years, excluded 2025, and repeated the main series without the masking correction. Analyses were performed in R 4.6.1 [24]. All code, the derived analytic dataset, the population file and the download manifest are deposited.

Results

Cohort characteristics

Between 2001 and 2025 there were 59,208 discharges with a primary diagnosis of F31 and 4,060 coded F30, the latter representing 6.4% of the combined set. Women accounted for 66.8% of F31 discharges. Among episodes with insurance recorded, 67.1% were covered by FONASA and 26.4% by ISAPRE; insurance was unknown or unspecified in a further share that is tabulated in the supplementary material rather than absorbed silently into the denominator. Median length of stay was 15 days (IQR 7–27) and changed little across the period, rising from 14 to 16 days between the two halves. In-hospital mortality was 0.09% (52 deaths). Unspecified bipolar disorder (F31.9) was the most frequent subtype throughout. Table 1 summarises these characteristics by period.

Table 1
Characteristics of hospital discharges with a primary diagnosis of bipolar disorder (ICD-10 F31), Chile, 2001–2025

Characteristic	2001–2013	2014–2025	Total
Discharges, n	27,242	31,966	59,208
Women, %	68.1	65.6	66.8
FONASA, %	64.8	68.9	67.1
ISAPRE, %	25.7	27.0	26.4
Other insurance, %	9.6	4.1	6.5
Length of stay, median days (IQR)	14 (6–26)	16 (7–28)	15 (7–27)
In-hospital mortality, %	0.11	0.07	0.09
Deaths, n	29	23	52
Aged 10–19, %	9.5	12.5	11.1
Aged 20–29, %	19.7	24.4	22.3
Aged 50 and over, %	25.3	25.7	25.5
Age-standardised rate, first → last year of period	10.0 → 11.6	11.7 → 16.1	—

Percentages for sex and insurance are computed among records with that variable recorded; the proportion masked by the data provider is reported in the supplementary material. Age-standardised rates (per 100,000, WHO world standard) begin in 2002, the first year covered by the official population projections. IQR interquartile range; FONASA national public insurer; ISAPRE private insurers.

The annual number of F31 discharges rose from 1,421 in 2001 to a maximum of 3,440 in 2024, with 3,279 in 2025. The 2025 file is not truncated: it contains more all-cause records (1,705,514) than any other year, so the small decline from 2024 reflects the data as published rather than incomplete consolidation.

Overall and sex-specific trends

The age-standardised rate rose from 10.0 per 100,000 (95% CI 9.3–10.7) in 2002 to 16.1 (15.6–16.7) in 2025, having peaked at 16.9 (16.3–17.5) in 2024. Averaged over the period the rate increased by 1.9% per year (1.4–2.5).

The trend was not monotonic. Joinpoint analysis placed knots at 2010, 2013 and 2021, defining four phases: a rise of 3.4% per year (1.7–5.1) to 2010; a decline of 6.2% per year (− 9.4 to − 2.9) between 2010 and 2013; near-stability at 1.4% per year (0.0–2.8) to 2021; and an acceleration of 6.6% per year (3.7–9.7) thereafter (Fig. 1, Table 2).

Table 2
Joinpoint regression of age-standardised and age-specific hospitalisation rates for bipolar disorder, Chile, 2002–2025

Series	Segment	APC, %	95% CI	AAPC, %	AAPC 95% CI
Total (F31)	2002–2010	3.4	1.7 to 5.1	1.9	1.4 to 2.5
	2010–2013	−6.2	−9.4 to − 2.9		
	2013–2021	1.4	0.0 to 2.8		
	2021–2025	6.6	3.7 to 9.7		
Women	2002–2009	4.5	2.3 to 6.7	2.0	1.4 to 2.6
	2009–2014	−4.9	−7.1 to − 2.7		
	2014–2021	1.9	0.2 to 3.7		
	2021–2025	6.9	3.5 to 10.4		
Men	2002–2022	1.0	0.5 to 1.5	1.8	1.4 to 2.3
	2022–2025	7.7	3.7 to 11.9		
Aged 0–9	2003–2007	30.4	−7.8 to 84.5	0.7	−5.3 to 7.1
	2007–2010	−32.3	−53.0 to − 2.5		
	2010–2017	22.4	6.6 to 40.5		
	2017–2024	−15.2	−26.2 to − 2.6		
Aged 10–19	2002–2013	3.4	1.2 to 5.6	4.9	4.1 to 5.8
	2013–2025	6.4	4.8 to 8.1		
Aged 20–29	2002–2010	3.4	1.5 to 5.4	3.7	3.1 to 4.2
	2010–2013	−11.5	−14.6 to − 8.2		
	2013–2025	8.0	7.1 to 9.0		
Aged 30–39	2002–2010	2.7	0.7 to 4.8	1.2	0.5 to 1.8
	2010–2013	−10.3	−14.1 to − 6.4		
	2013–2022	1.5	−0.1 to 3.0		
	2022–2025	8.8	3.7 to 14.1		
Aged 40–49	2002–2010	3.1	1.3 to 4.9	−0.0	−0.7 to 0.6
	2010–2014	−6.0	−8.7 to − 3.3		

Every segment of every series is reported. Knots were selected by the Bayesian information criterion. APC annual percent change; AAPC average annual percent change over the whole period.

Series	Segment	APC, %	95% CI	AAPC, %	AAPC 95% CI
	2014–2022	−2.7	−4.3 to − 1.1		
	2022–2025	7.6	2.2 to 13.2		
Aged 50–59	2002–2006	−1.7	−6.2 to 3.0	−0.1	−0.8 to 0.6
	2006–2009	15.0	10.6 to 19.6		
	2009–2022	−3.4	−4.1 to − 2.7		
	2022–2025	2.2	−1.9 to 6.5		
Aged 60–69	2002–2009	5.0	1.4 to 8.6	2.1	1.0 to 3.2
	2009–2022	−2.7	−4.0 to − 1.4		
	2022–2025	18.0	10.8 to 25.7		
Aged 70–79	2002–2019	−0.5	−1.7 to 0.8	−1.9	−3.0 to − 0.7
	2019–2022	−22.8	−29.5 to − 15.5		
	2022–2025	15.1	2.0 to 29.9		
Aged 80+	2002–2014	−0.1	−4.2 to 4.2	−6.6	−8.8 to − 4.4
	2014–2025	−13.3	−17.6 to − 8.8		

Every segment of every series is reported. Knots were selected by the Bayesian information criterion. APC annual percent change; AAPC average annual percent change over the whole period.

Rates in women exceeded those in men in every year, and the gap widened: the female-to-male ratio of standardised rates rose from 1.56 in 2002 to 1.91 in 2025. The two sexes followed different shapes. In women the series broke at 2009, 2014 and 2021; in men a single knot at 2022 separated two decades of slow increase (1.0% per year, 0.5–1.5) from a recent acceleration (7.7% per year, 3.7–11.9) that rests on four points.

Age redistribution

The most pronounced change was in who is admitted (Fig. 2). Rates in the 10–19 group rose from 5.96 to 17.20 per 100,000, with a single break at 2013 separating an increase of 3.4% per year (1.2–5.6) from one of 6.4% per year (4.8–8.1). Among those aged 20–29 the rate rose from 13.30 to 30.38 per 100,000, with the steepest phase, 8.0% per year (7.1–9.0), beginning in 2013.

Older groups moved in the opposite direction. Rates were essentially flat at 40–49 and 50–59 (AAPC − 0.0% and − 0.1%), declined at 70–79 (− 1.9% per year, − 3.0 to − 0.7) and fell substantially at 80 and over (− 6.6% per year, − 8.8 to − 4.5), the latter driven by a decline of 13.3% per year after 2014. Several of

these older series nonetheless show a common upturn after 2022, at 30–39, 40–49, 60–69 and 70–79, so part of the recent acceleration is not specific to young people.

In the multivariable model (Table S1), rates in women were 1.98 times those in men (95% CI 1.92–2.04). Relative to the 40–49 reference, rates were comparable at 20–29 and 30–39 and declined steadily thereafter, reaching 0.15 (0.13–0.18) at 80 and over. Period indicators describe average levels rather than slopes: relative to 2002–2009, the rate ratio was 1.09 (1.04–1.14) in 2010–2012, 1.00 (0.96–1.03) in 2013–2020 and 1.14 (1.10–1.19) in 2021–2025. The 2010–2012 period sits at a higher average level than the first even though the trend within it was falling, which is why a single linear term for calendar time would be uninterpretable here.

Sex differences across age groups

The marginal results, higher rates in women and rising rates in the young, are often combined into a claim about young women. Combining them does not establish it, so we tested it directly. The three-way interaction between age group, sex and period was significant ($F = 3.23$ on 8 df, $p = 0.001$), meaning the sex gap did change differently across age groups, but not in the direction that claim implies.

Between 2002–2013 and 2014–2025 the female-to-male rate ratio narrowed among young and middle-aged adults and widened among older ones. It fell from 1.68 to 1.41 at 20–29 and from 2.37 to 1.70 at 30–39, the latter the only clearly significant narrowing (ratio of ratios 0.72, 95% CI 0.61–0.84, $p < 0.001$). It rose from 1.44 to 2.33 at 70–79 (1.61, 1.12–2.31, $p = 0.013$) and from 2.01 to 2.57 at 60–69 (1.28, 0.97–1.68, $p = 0.089$). Among adolescents aged 10–19 the ratio moved from 1.70 to 2.00, but the change was imprecise and compatible with no difference (1.18, 0.87–1.60, $p = 0.30$). Table S4 gives the full set.

The widening of the overall sex gap therefore comes from older adults, among whom rates were falling in both sexes but falling faster in men, and not from adolescents or young adults, where the increase in admissions was steepest but affected both sexes more evenly.

Territorial disparity

Regional rates differed severalfold and the ranking was stable. Aysén had the highest standardised rate in both periods, rising from 28.0 to 38.8 per 100,000, while the lowest was Los Ríos in 2002–2013 (5.9) and Coquimbo in 2014–2025 (7.7); the extreme ratio was 4.74 and 5.02 respectively (Fig. 3a, Table S3).

Measures using the whole distribution give a different picture. The SII fell from 12.8 per 100,000 (95% CI 10.3–15.3) to 8.3 (4.0–12.7), and the RII from 2.90 to 1.89. Disparity between the extremes therefore persisted, and widened slightly, while inequality across the population as a whole narrowed. The two summaries are not contradictory: several intermediate regions converged toward the national average, and the residual gap is increasingly concentrated in the sparsely populated far south, whose small denominators also make its estimates the least precise (Fig. 3a).

The 2013 guarantee

Under the specification used in earlier work, a single pre-intervention slope fitted from 2002 with the population aged 10 and over, incorporation into the scheme was associated with a level reduction from 2014 (RR 0.83, 95% CI 0.73–0.95) and no change in slope (RR 1.01, 0.99–1.03); restricting to those aged 20 and over gave RR 0.81 (0.70–0.94).

That estimate is not robust to how the earlier trend is modelled. Allowing the slope break that the joinpoint analysis located at 2010, the level change attenuated to RR 0.94 (0.88–1.00) at 20 and over and RR 0.94 (0.88–1.01) at 10 and over, and restricting the pre-intervention window to 2010–2013 gave the same result (RR 0.94, 0.88–1.00; Table S2). Figure 3b, enlarged as Fig. S3, shows why. Rates had been falling steeply for three years before the guarantee took effect, and a model constrained to a single pre-intervention slope attributes that accumulated deviation to the intervention. Excluding the pandemic years or the 2013 transition year did not change this conclusion.

Sensitivity analyses

Combining F30 with F31 left the trend unchanged (AAPC + 1.2% versus + 1.0% for the log-linear comparison). Excluding the pandemic years slightly steepened it and excluding 2025 slightly flattened it, in neither case altering the interpretation. The complete-case analysis without the masking correction reproduced the trend from 2003 onward but, as expected, produced an implausibly low value for 2002 (supplementary material).

Discussion

Across a quarter of a century of national records, the composition of who is admitted for BD in Chile changed more than the total volume. Rates in adolescents roughly tripled and those in young adults more than doubled, while rates after age 70 fell; women were admitted at about twice the rate of men throughout, and that gap widened; and a fivefold difference between the extreme regions persisted although inequality across the whole population narrowed. The association previously reported between the 2013 guarantee and a step reduction in admissions did not survive a more careful specification of the pre-existing trend.

The age redistribution is the finding with the strongest claim on attention, and also the one most open to competing explanations. It is consistent with the international rise in mental health service use among adolescents and young adults described for depression, anxiety and self-harm [25–28], and with the global increase in BD incidence estimated for these age groups [4]. It is not, however, a universal pattern. The closest published counterpart, a Spanish national study of admissions for BD among 10–19-year-olds, found rates rising in the early 2000s and then declining moderately after 2009 [15], whereas the Chilean adolescent series breaks upward in 2013 and continues to climb through 2025. Two national registries, similar diagnostic codes and opposite recent directions are difficult to reconcile with a purely clinical account, and point instead toward service organisation and diagnostic practice.

Several mechanisms could produce the Chilean pattern. Psychopathology may genuinely be increasing in younger cohorts; detection may have improved and diagnosis moved earlier, with reduced stigma bringing younger patients into contact with services, a shift documented for paediatric BD internationally and still contested [29]; coding practice may have changed, with presentations formerly recorded as depressive episodes now recorded as bipolar; the guarantee itself eased access for people aged 15 and over from 2013, the year the adolescent series breaks; and admission thresholds may have fallen as adolescent inpatient capacity expanded, in a region where it started from a very low base [8].

Our data cannot adjudicate among these. One competing explanation deserves particular weight because it would reproduce the observed pattern exactly with no change in occurrence at all: an increase in the number of admissions per patient. Rates rose while median length of stay and in-hospital mortality stayed flat, which is what one would expect if the same population were being admitted more often rather than a larger population being admitted once. DEIS releases no patient identifier, so we could not estimate readmission frequency or test this. Any reading of these trends as evidence about the occurrence of BD in young Chileans is therefore premature.

The post-2021 acceleration invites a pandemic interpretation, given the documented deterioration in adolescent mental health during that period [25, 26] and the disruption and rebound of service use recorded elsewhere [30]. The age pattern complicates that reading: were the acceleration specific to youth mental health, it should concentrate in the youngest groups, yet the upturn after 2022 appears across several adult groups, including 60–69, where rates had been falling for over a decade. A general recovery of hospital activity after the 2020 contraction, when all-cause discharges fell from about 1.67 to 1.33 million, is at least as plausible for part of the rise.

The age-specific analysis also separates our findings from a common reading of this literature. Reports of rising service use in adolescents and young adults frequently emphasise young women, and it would have been natural to describe the Chilean increase the same way, since rates rose most in the young and are higher in women overall. The data do not support that description. Where the increase was steepest, at 10–29 years, the sex ratio was stable or narrowing; the widening of the overall gap came from people aged 60 and over, a group whose rates were falling. Marginal associations in age and in sex, each real on its own, would have produced a claim that a stratified analysis contradicts. The practical implication differs: services for young people face rising demand from both sexes, while the growing sex asymmetry sits in older adults. Why women are admitted at roughly twice the rate of men is not resolved by these data. Higher prevalence of bipolar II and of depression-predominant presentations in women, both associated with more frequent service contact, is one explanation consistent with the literature [1, 2]; a differential in readmission frequency by sex would be another, and the episode-level structure of the data cannot distinguish them.

The territorial finding needs to be stated more carefully than the extreme ratio alone allows. Aysén and Magallanes sustained high rates throughout, which most likely reflects thinner outpatient alternatives, fewer community mental health teams and greater distances rather than higher occurrence; regions with

little capacity to manage crises in the community admit more of them. The same gradient has been described for self-harm hospitalisation across levels of rurality in Canada [31], and regional disparities in depression and healthy life expectancy have been reported within Chile itself [9]. Chile reduced its psychiatric hospital beds substantially while expanding community services over this period [11], and bed provision across Latin America remains both low and unevenly distributed [8], so supply is a plausible driver of the residual gap. That interpretation is untested here, since we did not link the discharge data to regional indicators of bed supply or outpatient capacity, which is a natural next step. The divergence between a stable extreme ratio and a falling RII carries a practical message: the average region has moved toward the national mean, and what remains is a concentrated gap in the far south affecting small populations, which is a different policy problem from generalised territorial inequality and calls for different instruments.

The policy result illustrates how sensitive inference from segmented regression is to assumptions about what would have happened otherwise. Reported under a single pre-intervention slope, the guarantee appears to have cut admissions by about 17%. Once the decline already under way since 2010 is modelled, the estimate moves to a 6% reduction whose confidence interval includes no effect. Nothing about the data changed; only the counterfactual did. We cannot say from these data whether the guarantee reduced admissions, and the earlier interpretation, that intermediate-severity episodes shifted toward guaranteed outpatient care, is not supported by this analysis. What drove the 2010–2013 decline remains open.

This has a bearing beyond the present study. Segmented regression is the standard tool for evaluating Chilean health guarantees and mental health plans [11] and is widely used for service-use series in the region [12], yet long national series will often contain trend breaks unrelated to the policy under study. Where a joinpoint analysis and an interrupted time series are applied to the same data, the knots identified by the former should inform the pre-intervention specification of the latter rather than be reported alongside it as an independent result. Brazil's sustained decline over a period of community expansion [13], and the contrasting English and Czech trajectories [5, 6], show that secular trends in this outcome can be large enough to swamp the effect of any single measure.

Strengths and limitations

The study covers all public and private discharges nationally over 25 years, uses official denominators with age standardisation, reports every joinpoint segment rather than a selection, quantifies territorial inequality with measures that use the full distribution, and makes the entire pipeline reproducible from the original open data.

Its limitations are mostly inherent to the source. The registry captures episodes rather than individuals, so rates describe hospitalisations rather than incidence or prevalence, and readmissions cannot be identified. Diagnoses rely on administrative coding, subject to error and to drift over time, and cannot be validated against clinical records. Age is available only in grouped form, which limits the precision of

standardisation and prevents both the isolation of the guarantee-eligible population at age 15 and the 10–14 versus 15–19 comparison that would have provided an internal control. The quality of region of residence is uneven for Arica y Parinacota, whose all-cause counts fluctuate implausibly across the whole series rather than only before its creation in 2007, suggesting intermittent assignment of its records to Tarapacá; the two regions are better interpreted jointly. Completeness of sex, age and region is high overall but poor in 2002, and our correction assumes masking is unrelated to those characteristics within a year, an assumption we cannot verify. Insurance is recorded for most but not all episodes, so the FONASA and ISAPRE shares describe the recorded subset and are not insurance-specific rates. Finally, the policy analysis is ecological and, as shown, sensitive to specification; it supports no individual-level causal inference.

Conclusion

Hospital admission for BD in Chile became both more frequent and younger between 2002 and 2025, with rates in adolescents and young adults rising while those in older adults fell, and with women admitted at roughly twice the rate of men throughout. These are changes in the use of inpatient care, and the data cannot establish whether the occurrence of the disorder itself has changed. Planning should nonetheless follow the demand: services for adolescents and young adults, and the transition between paediatric and adult care, now carry a share of admissions that they did not two decades ago. The territorial pattern calls for instruments targeted at the far south rather than general measures, since inequality across the rest of the country has narrowed. The evidence does not support attributing the 2013 change to the Explicit Health Guarantees, and evaluations of that policy will need designs that do not rest on extrapolating a single pre-intervention slope.

Declarations

Competing interests

The authors have no relevant financial or non-financial interests to disclose.

Funding

No funding was received for conducting this study.

Ethics approval

The study used exclusively anonymised, publicly available administrative databases published as open data by the Department of Health Statistics and Information of the Chilean Ministry of Health, containing no personal identifiers. Under Chilean Law No. 19,628 on the protection of personal data, secondary analysis of such dissociated public records does not require ethics committee approval or informed

consent, and none was therefore sought. The study was conducted in accordance with the Declaration of Helsinki.

Consent to participate

Not applicable. The study analyses anonymised administrative records released as open data and involves no identifiable individual data, so no individual consent could be or was sought.

Consent to publish

Not applicable. The manuscript contains no data or images relating to an identifiable individual.

Data availability

The original hospital discharge records are publicly available from the DEIS Open Data repository (https://repositoriodeis.minsal.cl/DatosAbiertos/EGRESOS/) and the population projections from the National Statistics Institute of Chile (https://www.ine.gob.cl), both accessed on 9 August 2026. The derived analytic dataset, the harmonised population file, the complete analysis code and a manifest recording the size, modification date and SHA-256 checksum of every source file are deposited in the Open Science Framework at https://doi.org/10.17605/OSF.IO/Y67F8.

Authors' contributions

G.A.-O. and F.A.-G. conceived and designed the study jointly and agreed the analysis plan. F.A.-G. retrieved and harmonised the open administrative data, wrote the analysis code, and prepared the tables and figures. G.A.-O. conducted the literature review and led the clinical and health-policy interpretation of the findings. Both authors checked the analyses against the source data, wrote the main manuscript text, revised successive drafts, and approved the submitted version. G.A.-O. and F.A.-G. contributed equally to this work.

Use of generative artificial intelligence

During the preparation of this work the authors used Claude Sonnet 5 (Anthropic) to assist with the drafting of analysis code and with language editing. After using this tool, the authors reviewed and edited the content as needed and take full responsibility for the content of the published article. The tool is not an author and is not credited as one.

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References

1. Vieta E, Berk M, Schulze TG, Carvalho AF, Suppes T, Calabrese JR, Gao K, Miskowiak KW, Grande I (2018) Bipolar disorders. Nat Rev Dis Primers 4:18008. https://doi.org/10.1038/nrdp.2018.8
2. McIntyre RS, Berk M, Brietzke E, Goldstein BI, López-Jaramillo C, Kessing LV, Malhi GS, Nierenberg AA, Rosenblat JD, Majeed A (2020) Bipolar disorders. Lancet 396:1841–1856. https://doi.org/10.1016/S0140-6736(20)31544-0
3. Oliva V, Fico G, De Prisco M, Gonda X, Rosa A, Vieta E (2025) Bipolar disorders: an update on critical aspects. Lancet Reg Health Eur 48:101135. https://doi.org/10.1016/j.lanepe.2024.101135
4. Zhong Y, Chen Y, Su X, Wang M, Li Q, Shao Z, Sun L (2024) Global, regional and national burdens of bipolar disorders in adolescents and young adults: a trend analysis from 1990 to 2019. Gen Psychiatr 37:e101255. https://doi.org/10.1136/gpsych-2023-101255
5. Degli Esposti M, Ziauddeen H, Bowes L, Reeves A, Chekroud AM, Humphreys DK, Ford T (2022) Trends in inpatient care for psychiatric disorders in NHS hospitals across England, 1998/99–2019/20: an observational time series analysis. Soc Psychiatry Psychiatr Epidemiol 57:993–1006. https://doi.org/10.1007/s00127-021-02215-5
6. Potočár L, Winkler P, Mohr P, Formánek T (2024) Temporal trends in inpatient care use for adult mental disorders in Czechia: a nationwide register-based study from 1994 to 2015. Soc Psychiatry Psychiatr Epidemiol 59:1825–1837. https://doi.org/10.1007/s00127-024-02691-5
7. Saz-Parkinson Z, Medel A, Cediel-García P, Castellote J, Bouza C, Amate JM (2011) Trends on schizophrenia admissions during the deinstitutionalisation process in Spain (1980–2004). Soc Psychiatry Psychiatr Epidemiol 46:1095–1101. https://doi.org/10.1007/s00127-010-0289-9
8. Mundt AP, Delhey Langerfeldt S, Martínez P, Irarrázaval M, Grasso L, Rivera-Arroyo G, Capistrano A, Trujillo N, Golcher F, Benavides Salcedo A (2024) Types of psychiatric beds and mental health services in 16 Latin American countries, 1990–2020. Psychiatr Serv 75:48–54. https://doi.org/10.1176/appi.ps.20220590
9. Rueda-Salazar S, Miranda-Castillo C, Araya AX (2025) Regional and gender disparities in depression and late life expectancy in Chile. Soc Psychiatry Psychiatr Epidemiol 60:2359–2376. https://doi.org/10.1007/s00127-025-02883-7
10. Ministry of Health of Chile (2013) GES clinical practice guideline: bipolar disorder in people aged 15 and over. MINSAL, Santiago
11. Mundt AP, Martínez P, Jaque S, Irarrázaval M (2022) The effects of national mental health plans on mental health services development in Chile: retrospective interrupted time series analyses of national databases between 1990 and 2017. Int J Ment Health Syst 16:5. https://doi.org/10.1186/s13033-022-00519-w

12. Villarreal-Zegarra D, Reátegui-Rivera CM, Otazú-Alfaro S, Yantas-Alcantara G, Soto-Becerra P, Melendez-Torres GJ (2023) Estimated impact of the COVID-19 pandemic on the prevalence and treatment of depressive symptoms in Peru: an interrupted time series analysis in 2014–2021. Soc Psychiatry Psychiatr Epidemiol 58:1375–1385. https://doi.org/10.1007/s00127-023-02446-8
13. Resende BL, Fonseca Guimarães JS, Amâncio Caetano AL, Matos De Oliveira AL, Rezende Souza B, Oliveira AG (2026) Nationwide decline in psychiatric hospitalizations and costs in Brazil, 2008–2022: a retrospective descriptive study evidencing reform-driven community mental health gains. Lancet Reg Health Am 56:101425. https://doi.org/10.1016/j.lana.2026.101425
14. Rodríguez-Lorenzana A, Coral-Almeida M, Carrington SJ, Torres-Tapia M, Álvarez-Mejía D, Santana M, Mascialino G (2025) Trends, geographic distribution, and disease burden of bipolar disorder in Ecuador (2011–2021): an analysis of hospital discharge data. PLoS One 20:e0320321. https://doi.org/10.1371/journal.pone.0320321
15. López-Cuadrado T, Susser E, Martínez-Alés G (2024) Recent trends in hospital admission due to bipolar disorder in 10–19-year-olds in Spain: a nationwide population-based study. Bipolar Disord 26:801–809. https://doi.org/10.1111/bdi.13500
16. Department of Health Statistics and Information (DEIS) (2026) Hospital discharge open data, 2001–2025. Ministry of Health of Chile, Santiago. https://repositoriodeis.minsal.cl/DatosAbiertos/EGRESOS/. Accessed 9 August 2026
17. Benchimol EI, Smeeth L, Guttmann A, Harron K, Moher D, Petersen I, Sørensen HT, von Elm E, Langan SM (2015) The REporting of studies Conducted using Observational Routinely-collected health Data (RECORD) statement. PLoS Med 12:e1001885. https://doi.org/10.1371/journal.pmed.1001885
18. National Statistics Institute of Chile (INE) (2018) Estimaciones y proyecciones de la población de Chile 2002–2035, regiones, base 2017. INE, Santiago. https://www.ine.gob.cl. Accessed 9 August 2026
19. Ahmad OB, Boschi-Pinto C, Lopez AD, Murray CJL, Lozano R, Inoue M (2001) Age standardization of rates: a new WHO standard. GPE discussion paper series no. 31. World Health Organization, Geneva
20. Fay MP, Feuer EJ (1997) Confidence intervals for directly standardized rates: a method based on the gamma distribution. Stat Med 16:791–801
21. Morthorst BR, Heinrichsen M, Erlangsen A (2025) Joinpoint analyses of rates on hospital-recorded deliberate self-harm: an update on Danish national trends. Soc Psychiatry Psychiatr Epidemiol 60:727–736. https://doi.org/10.1007/s00127-024-02795-y
22. Kim HJ, Fay MP, Feuer EJ, Midthune DN (2000) Permutation tests for joinpoint regression with applications to cancer rates. Stat Med 19:335–351
23. Wagner AK, Soumerai SB, Zhang F, Ross-Degnan D (2002) Segmented regression analysis of interrupted time series studies in medication use research. J Clin Pharm Ther 27:299–309
24. R Core Team (2026) R: a language and environment for statistical computing. R Foundation for Statistical Computing, Vienna. https://www.R-project.org/. Accessed 9 August 2026

25. Santomauro DF, Mantilla Herrera AM, Shadid J, Zheng P, Ashbaugh C, Pigott DM, et al (2021) Global prevalence and burden of depressive and anxiety disorders in 204 countries and territories in 2020 due to the COVID-19 pandemic. Lancet 398:1700–1712. https://doi.org/10.1016/S0140-6736(21)02143-7
26. Racine N, McArthur BA, Cooke JE, Eirich R, Zhu J, Madigan S (2021) Global prevalence of depressive and anxiety symptoms in children and adolescents during COVID-19: a meta-analysis. JAMA Pediatr 175:1142–1150. https://doi.org/10.1001/jamapediatrics.2021.2482
27. Cybulski L, Ashcroft DM, Carr MJ, Garg S, Chew-Graham CA, Kapur N, Webb RT (2021) Temporal trends in annual incidence rates for psychiatric disorders and self-harm among children and adolescents in the UK, 2003–2018. BMC Psychiatry 21:229. https://doi.org/10.1186/s12888-021-03235-w
28. Parodi KB, Holt MK, Green JG, Porche MV, Koenig B, Xuan Z (2022) Time trends and disparities in anxiety among adolescents, 2012–2018. Soc Psychiatry Psychiatr Epidemiol 57:127–137. https://doi.org/10.1007/s00127-021-02122-9
29. Goldstein BI, Birmaher B, Carlson GA, DelBello MP, Findling RL, Fristad M, Kowatch RA, Miklowitz DJ, Nery FG, Pérez-Algorta G (2017) The International Society for Bipolar Disorders Task Force report on pediatric bipolar disorder: knowledge to date and directions for future research. Bipolar Disord 19:524–543. https://doi.org/10.1111/bdi.12556
30. Joo SW, Kim H, Jo YT, Ahn S, Choi YJ, Choi W, Lee J (2023) Impact of the early phase of the COVID-19 pandemic on the use of mental health services in South Korea: a nationwide, health insurance data-based study. Soc Psychiatry Psychiatr Epidemiol 58:441–452. https://doi.org/10.1007/s00127-022-02382-z
31. Mahinpey N, Pollock NJ, Liu L, Contreras G, Thompson W (2023) Self-harm and rurality in Canada: an analysis of hospitalization data from 2015 to 2019. Soc Psychiatry Psychiatr Epidemiol 58:1161–1170. https://doi.org/10.1007/s00127-023-02463-7

Figures

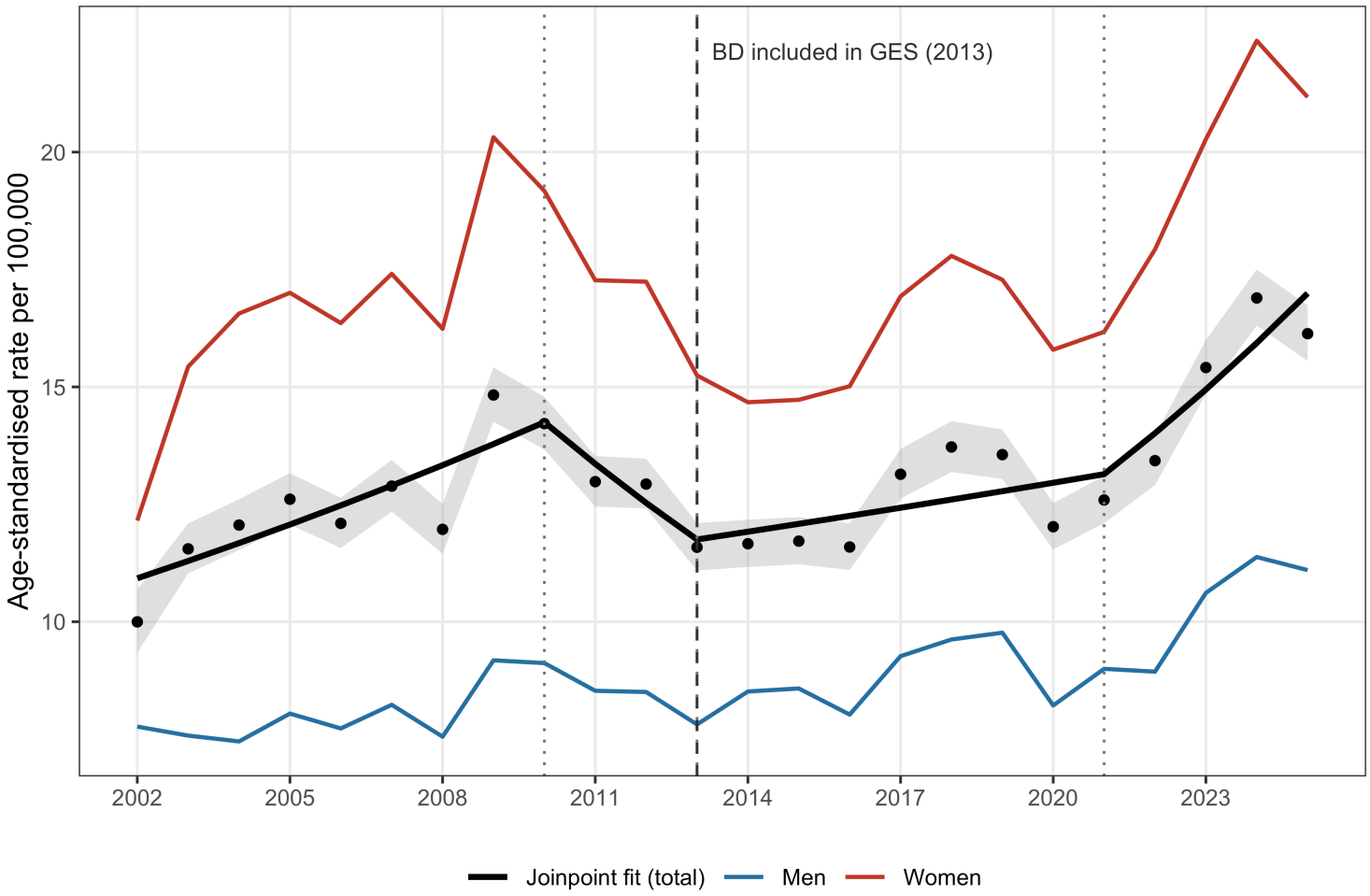


Figure 1

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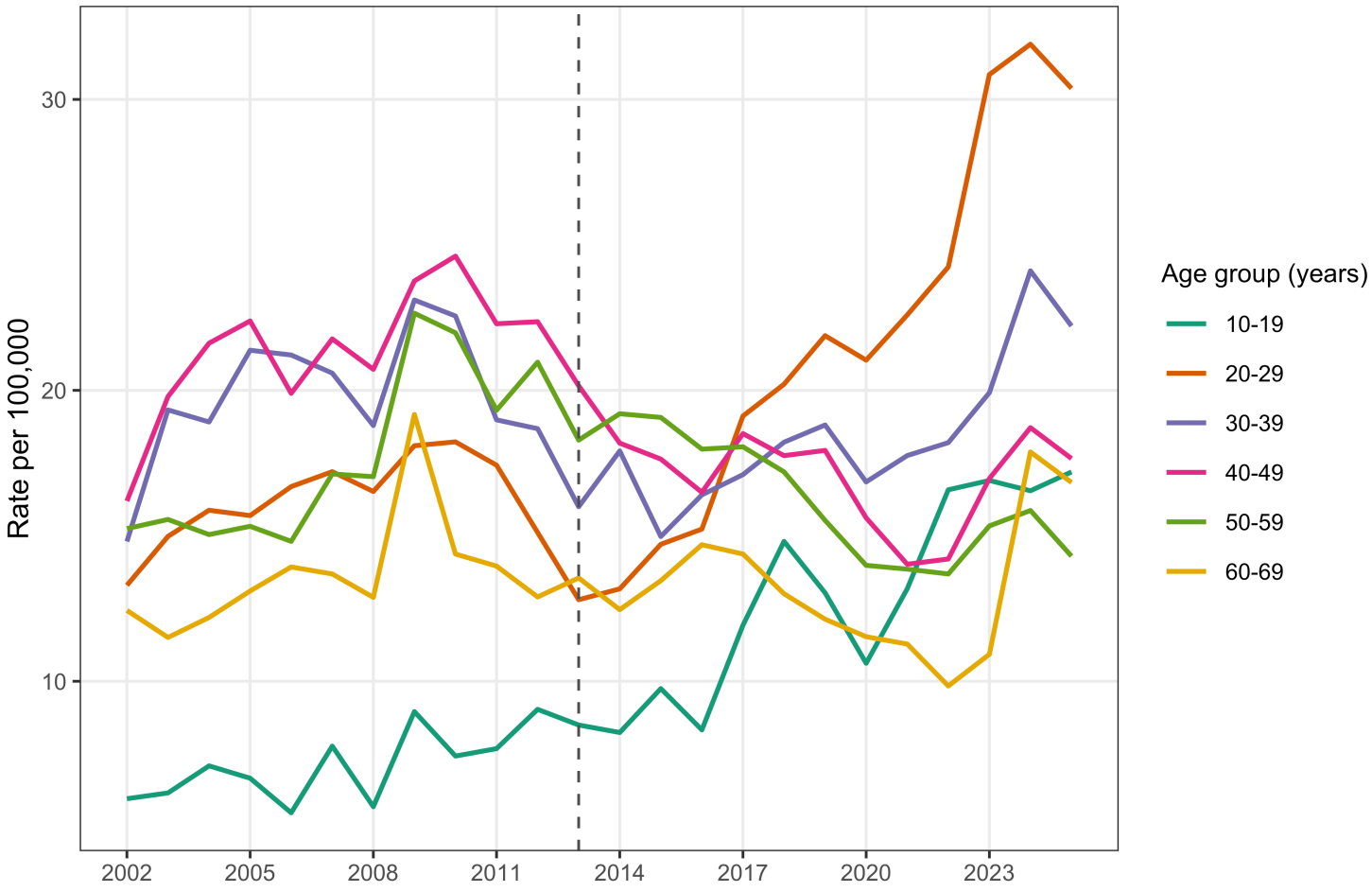


Figure 2

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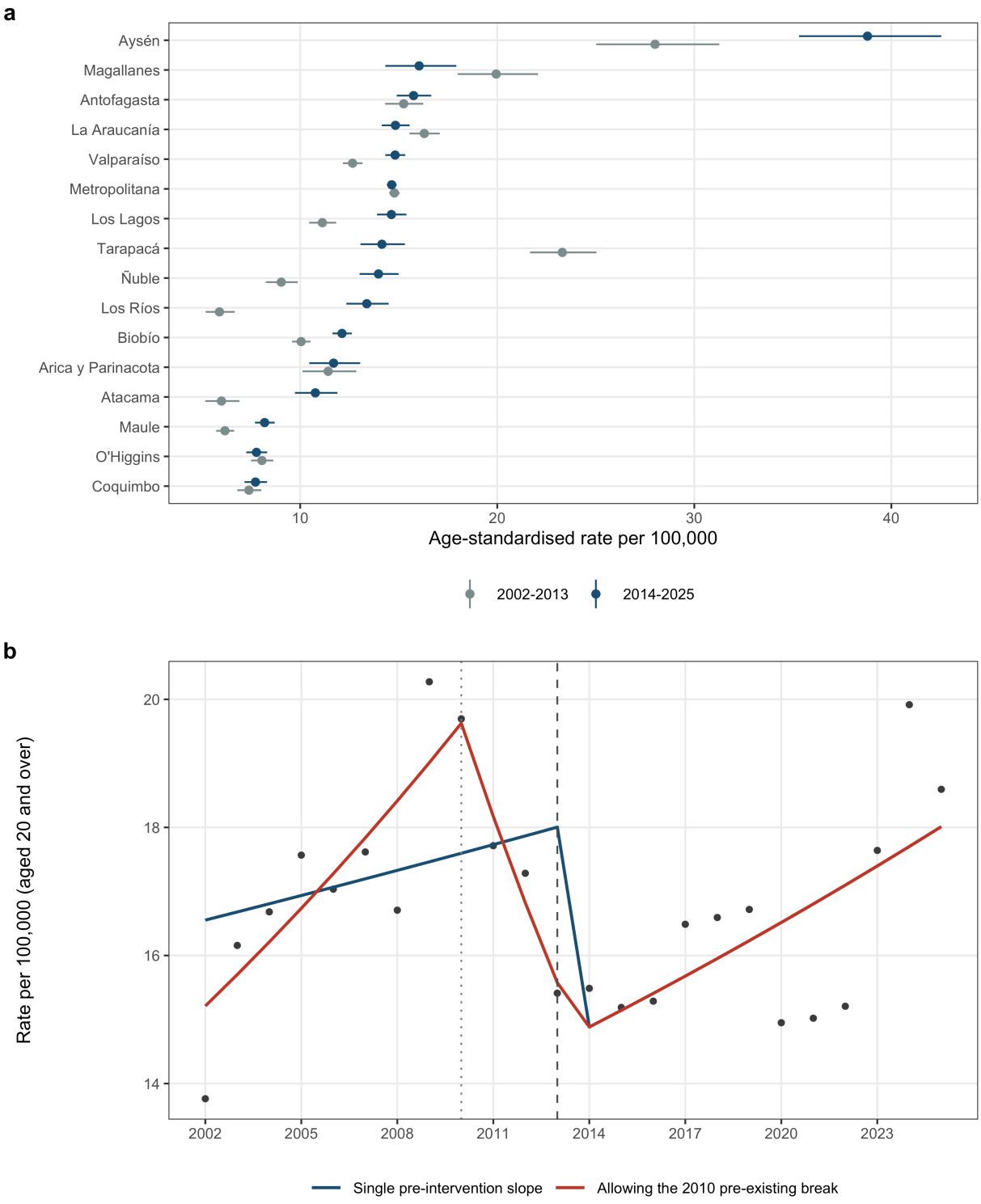


Figure 3

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