Economics Working Paper Series
Faculty of Business, Economics and Law
Auckland University of Technology

Gender Identity and Recorded Offending: Evidence from Linked Population Data

Maxine J. Lee and Alexander Plum

2026/05

Gender Identity and Recorded Offending: Evidence from Linked Population Data*

Maxine J. Lee and Alexander Plum†

October 2026

Abstract

Population evidence on gender identity and offending remains scarce. We link New Zealand's 2023 Census to administrative records for 170,775 New Zealand-born individuals from the 1993–1998 birth cohorts. Separate measures of sex at birth and gender distinguish six groups: cisgender women, cisgender men, transgender women, transgender men, and nonbinary individuals grouped by sex assigned at birth. Their histories of selected Police-recorded interpersonal and victim-related offenses are examined over a common ten-year window at ages 17–26. Adjusting for ethnicity, parental income, parental sentencing history, and year-month of birth, all four gender-minority groups have significantly lower odds of a qualifying record than cisgender men. For transgender women, the adjusted probability is 5.1%, compared with 12.2% for cisgender men and 6.0% for cisgender women. The corresponding differences are −7.2 and −1.0 percentage points; only the former is statistically significant. Across the main analyses of offense type, court action, victim characteristics, and location, transgender women consistently have lower adjusted probabilities than cisgender men, while differences from cisgender women are not statistically significant. These findings document variation in recorded offending obscured by conventional binary classifications.

JEL Classification: K42, J16, J15

Keywords: Gender identity; Police-recorded offending; Linked administrative data.

Stats NZ IDI Disclaimer: These results are not official statistics. They have been created for research purposes from the Integrated Data Infrastructure (IDI) which is carefully managed by Stats NZ. For more information about the IDI please visit `https://www.stats.govt.nz/integrated-data/`. *Disclaimer for Census 2023 data:* Access to the data used in this study was provided by Stats NZ under conditions designed to give effect to the security and confidentiality provisions of the Data and Statistics Act 2022. The results presented in this study are the work of the author, not Stats NZ or individual data suppliers. *Disclaimer for Inland Revenue tax data:* The results are based in part on tax data supplied by Inland Revenue to Stats NZ under the Tax Administration Act 1994 for statistical purposes. Any discussion of data limitations or weaknesses is in the context of using the IDI for statistical purposes, and is not related to the data's ability to support Inland Revenue's core operational requirements.

†Lee is Associate Professor of Economics at San Francisco State University, mclee@sfsu.edu. Plum is Associate Professor (Research) and Associate Director of the New Zealand Policy Research Institute, Auckland University of Technology (AUT), alexander.plum@aut.ac.nz. We are grateful to Michael Martell, Lisa Tompson, Karinna Saxby and seminar participants at the CSQIEP Virtual Seminar on LGBTQ+ Economics for helpful comments. All remaining errors are our own.

1 Introduction

Criminal offending imposes substantial social costs that extend beyond direct financial losses, including adverse effects on mental health and well-being (Cornaglia et al., 2014; Adams et al., forthcoming). Offending rates also exhibit pronounced age and sex differences: they typically peak in late adolescence and early adulthood (Bell et al., 2018; Dasgupta et al., 2022) and are substantially higher among men than women (Campaniello and Gavrilova, 2018). Research into the origins of offending highlights the roles of intergenerational transmission (Hjalmarsson and Lindquist, 2013), early family experiences and behavioral development (Currie and Tekin, 2012; Heckman et al., 2013), neighborhood and peer environments (Damm and Dustmann, 2014; Deming, 2011), and economic conditions during the transition to adulthood (Bell et al., 2018). Using Swedish register data, Hederos Eriksson et al. (2016) document substantial sibling similarities in criminal convictions and incarceration, with parental criminality and family structure accounting for more of these similarities than parental income and education or neighborhood characteristics.

This paper contributes to this literature by examining Police-recorded offending among gender-minority people, including transgender and non-binary individuals. Recent evidence from the United States documents a marked increase in the share of young adults identifying as transgender or gender nonconforming (Twenge et al., 2025). Existing research also documents substantial economic disadvantages among gender-minority people (Carpenter et al., 2020, 2022, 2024, 2025), alongside higher rates of criminal victimization (Flores et al., 2021) and poorer mental health (Saxby et al., 2025) relative to cisgender peers. Related research on sexual minorities documents disparities in health and social relationships, with substantial variation across sexual-orientation groups and institutional contexts (Berlingieri and Kovacic, 2025). This heterogeneity motivates distinguishing population groups rather than treating sexual and gender minorities as a single category. We contribute by documenting Police-recorded offending across groups defined jointly by sex at birth and gender identity.

We focus on interpersonal and other victim-related offenses because of their direct harm to victims and their broader implications for community safety. The consequences of such offenses can extend well beyond the immediate incident, affecting victims' economic opportunities, health, and family members. For example, sexual assault can reduce victims' employment and earnings and worsen their mental health for at least a decade, while also adversely affecting their parents' mental health (Adams et al., forthcoming). Related research documents substantial economic losses associated with abusive relationships (Adams et al., 2024) and persistent labor market consequences of violence between colleagues (Adams-Prassl et al., 2024). Understanding varying patterns of offending by gender identity can therefore inform policies to reduce harm and improve community

safety.

The relationship between gender identity and recorded offending is theoretically ambiguous. Restricted labor market opportunities may reduce the opportunity cost of offending, whereas anticipated harms associated with detention or imprisonment may increase its perceived costs. Differences in reporting, detection, and attribution to an identified person may also influence entry into Police records. Existing evidence provides limited guidance. For example, Dhejne et al. (2011) examine criminal convictions in a Swedish cohort of people who underwent surgical sex reassignment during 1973–2003. Such treatment-defined samples do not represent the broader population identified through separate measures of sex at birth and gender. Our self-identified sample allows us to examine recorded offending without requiring a history of gender-affirming medical treatment.

We address this evidence gap by linking the 2023 New Zealand Census to administrative records within Statistics New Zealand's Integrated Data Infrastructure (IDI). The Census asks sex at birth and gender in separate questions, allowing us to distinguish cisgender men, cisgender women, transgender men, transgender women, and non-binary people assigned male at birth (AMAB) or assigned female at birth (AFAB). We link these responses to Police records to examine recorded offending histories without relying on respondents' recall or disclosure of offending. Further linkages to birth registrations, parental income, and parental sentencing records provide measures of family background preceding the outcome window. Together, these sources allow us to document differences across six gender-identity groups and assess whether they persist after standardization for observed demographic and family-background characteristics.

We focus on the *extensive margin*: whether an individual has at least one qualifying Police offender record within the observation window. This outcome reflects reporting, detection, and attribution to an identified person as well as underlying behavior; our analysis does not separate these processes. It also does not measure the frequency, persistence, or cumulative severity of offending. The estimates therefore describe differences in the probability of a qualifying record, rather than differences in the total volume of offending or in all forms of contact with law enforcement.

Our analysis follows 170,775 New Zealand-born individuals from the 1993–1998 birth cohorts over a common ten-year age window, from their 17th birthday to immediately before their 27th birthday. Because the probability of ever having a recorded offense depends on both age and the length of observation, comparisons based on differently aged populations or unequal follow-up can conflate group differences with differences in accumulated exposure. Our design holds the age range and nominal observation length constant across all six gender groups, covering late adolescence and early adulthood, when offending is particularly concentrated. We additionally restrict extended overseas absences to limit differences in exposure to New Zealand administrative recording. Using

this sample, we estimate the adjusted probabilities of having a police record for selected offenses for each of the six gender groups and the odds ratios comparing the five gender groups, including four gender-minority groups and cisgender women, to cisgender men. Alongside the comparisons, we report the size of the gap in adjusted probabilities between transgender women and both cisgender men and cisgender women and measure the statistical significance of the gap. These additional comparisons reflect the prominence of transgender women in political discussions linking gender identity to public safety.

The selected offenses comprise assault, harm or endangerment of persons, homicide, robbery, blackmail and extortion, sexual offenses, and theft. These categories capture interpersonal and material harms with direct implications for victims and community safety.[1] We examine any qualifying record, assault-related offending, serious offending, and Police-recorded court action. We also distinguish records by victim sex, offender–victim relationship, and offense location. We report average adjusted probabilities and direct probability differences, supplemented by adjusted odds ratios.

Our central finding is that all four gender-minority groups have lower adjusted probabilities of any qualifying Police record than cisgender men. Standardizing groups to a common distribution of observed characteristics, the estimated probability is 12.2% among cisgender men, compared with 6.0% among cisgender women, 8.2% among transgender men, 5.1% among transgender women, 6.4% among non-binary people AMAB, and 4.6% among non-binary people AFAB. Each gender-minority group's adjusted odds are significantly below those of cisgender men at the 5% level. These findings show that the conventional male–female distinction conceals substantial variation in recorded offending across groups defined by both sex and gender.

For transgender women, the adjusted probability is 7.2 percentage points below that of cisgender men and 1.0 percentage point below that of cisgender women; the latter difference is not statistically significant. In the main analyses disaggregated by offense type, court action, victim characteristics, and location, transgender women also have significantly lower adjusted probabilities than cisgender men. None of the corresponding probability differences from cisgender women is statistically significant at the 5% level. These findings do not establish equivalence, particularly for less frequent outcomes and smaller population groups.

The contribution extends beyond the New Zealand setting by addressing an evidence gap at the intersection of research on gender differences in offending and the circumstances of gender-minority populations. Separate measures of sex at birth and gender, linked to administrative offending histories, allow us to examine population differences that conventional binary classifications cannot reveal. The findings provide an empirical reference point for research and public discussion concerning gender identity and recorded

[1]Research on public perceptions links the seriousness of the crime to perceived harmfulness and moral wrongfulness (Adriaenssen et al., 2020; Treshenkov et al., 2026).

offending. Although the estimated levels and gaps need not generalize to countries with different social and criminal-justice institutions, the study provides a framework for comparable analyses as suitable data become available elsewhere.

The remainder of the paper is organized as follows. Section 2 describes the linked data, sample construction, and descriptive statistics. Section 3 presents the empirical approach and explains the interpretation of adjusted probabilities and odds ratios. Section 4 reports the main results and supplementary analyses, and Section 5 concludes. Additional tables are provided in the Online Appendix.

2 Data and descriptive statistics

2.1 Data sources and linkage

This study uses the Integrated Data Infrastructure (IDI), a longitudinal research database maintained by Statistics New Zealand (Stats NZ). The IDI links individual-level data from government agencies, Stats NZ surveys, and other data providers. Linkage makes it possible to study outcomes recorded in different administrative systems without researchers having access to direct personal identifiers. Access is limited to approved research projects and approved researchers working in a secure Stats NZ Data Lab environment, and all results are subject to confidentiality checking before release. Our population is drawn from the 2023 New Zealand Census, which collected sex at birth and gender in separate questions. Figure 1 shows that we link 2023 Census records to Department of Internal Affairs (DIA) birth registrations, Police offender and victimization records, the IDI total-income module, and Ministry of Justice (MoJ) court-charge records. We also use the residential-address and overseas-spell tables prepared by Stats NZ. These linkages provide individual offending histories over a common age window and selected measures of family background before that window begins.

2.2 Study population and gender identity

The study population comprises New Zealand-born individuals from the 1993–1998 birth cohorts who provided self-responses on sex at birth and gender in the 2023 Census and can be linked to a DIA birth registration. The birth registrations identify recorded parents, enabling us to construct family-background measures.

Gender identity is classified using the 2023 Census, which asked sex at birth and gender of each respondent in separate questions for the first time (see Figure 2). The sex-at-birth response options were 'male' and 'female'; the gender options were 'male', 'female', and 'another gender', with a free-text field for the latter.[2] We classify respondents reporting

[2]We recognize the limitations of having only two options for sex at birth and the potential issue with

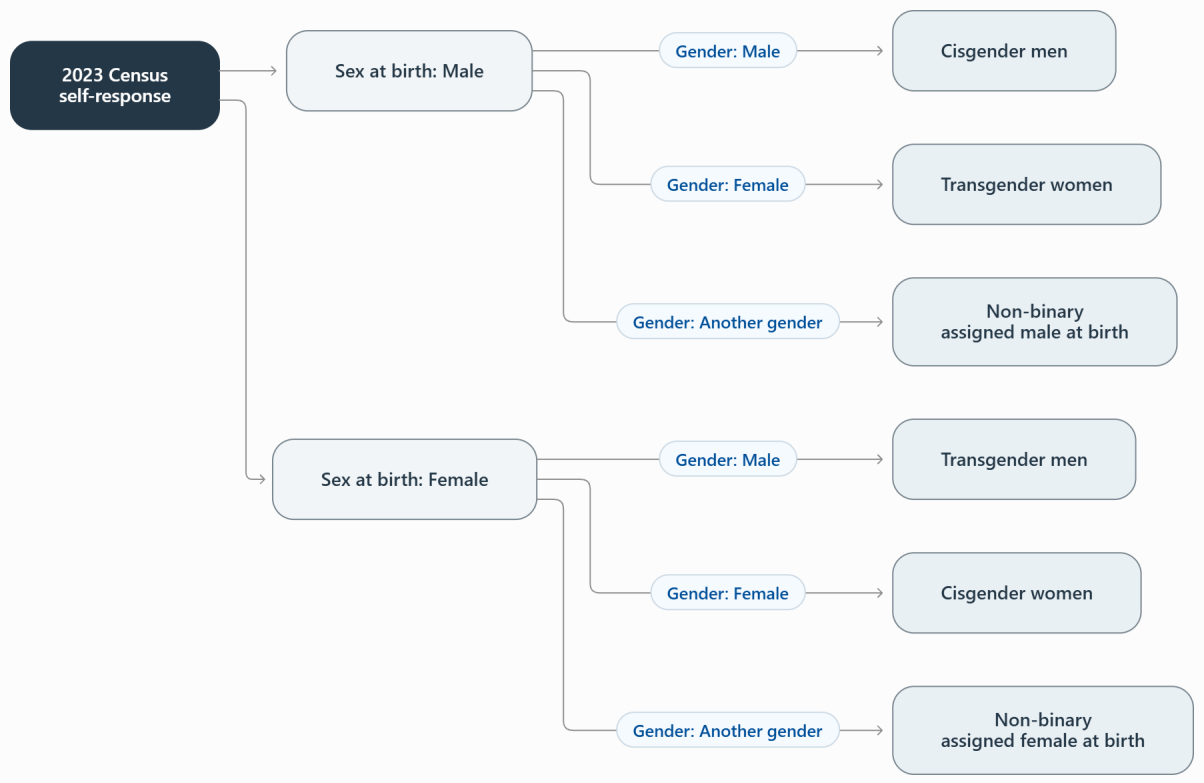
Figure 1: Data sources and construction of the linked analysis dataset



Notes: The diagram summarizes the sources for the principal measures; additional sample restrictions are described in Section 2. DIA denotes the Department of Internal Affairs, MoJ denotes the Ministry of Justice. The main offending window covers ages 17–26; outcomes involving victim sex are measured at ages 21–26. The figure presents sources used in the main analysis. Supplementary analyses add pharmaceutical dispensing and education records.

male sex at birth and male gender as cisgender men, and those reporting female sex at birth and female gender as cisgender women. Respondents reporting female sex at birth and male gender are classified as transgender men; those reporting male sex at birth and female gender are classified as transgender women. Stats NZ does not provide the text responses for gender and instead and instead aggregates entries into a single category 'another gender', which are then grouped as non-binary and distinguished by sex at birth. This is an analytical grouping and does not imply that every respondent in the category personally uses the term non-binary. We distinguish non-binary people by sex at birth to allow for heterogeneity that would be obscured by combining these groups. (Carpenter et al., 2020, 2025).

Figure 2: Classification of gender identity using 2023 Census responses.



Notes: Groups are derived jointly from self-reported sex at birth and gender. 'Another gender' includes free-text responses classified by Stats NZ in that category and is labeled non-binary in this study.

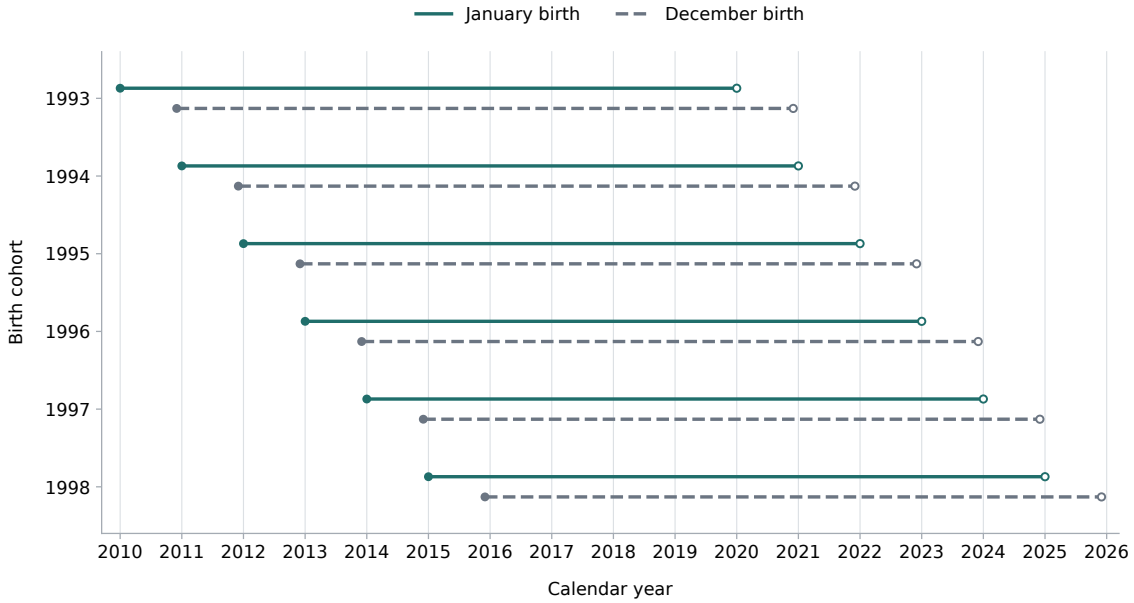
2.3 Observation window and residence restrictions

The choices of birth cohorts and the window of observation balance the breadth of cohort coverage against the length of follow-up. Police offender records are available in the IDI from June 2009 through December 2025. Requiring complete observation over a

describing gender with the terms 'female' and 'male', rather than 'women' and 'men'. Although the text specifies 'male' and 'female', we use 'women' and 'men' to refer to individuals' gender in the remainder of this article.

common age window therefore creates a trade-off: extending follow-up to older ages reduces the number of recent birth cohorts that can be included, while including more recent cohorts requires ending follow-up at a younger age. We select ages 17–26 to cover late adolescence and the early twenties, when offending is particularly concentrated, together with subsequent years of early adulthood (Dasgupta et al., 2022).

Figure 3: Birth cohorts and outcome observation windows



Notes: The sample comprises individuals born in 1993–1998. Outcomes are measured from the 17th birthday (filled circles) up to, but excluding, the 27th birthday (open circles), giving each individual a ten-year observation window. Solid and dashed lines illustrate January and December births, respectively; observation windows for other birth months lie between these examples. Ages are calculated using birth and offense months. Separate analyses distinguish ages 17–21 and 22–26, each covering five years.

Given this age window, 1993–1998 is the widest range of complete calendar-year birth cohorts with full administrative coverage. All individuals born in 1993 turn 17 after the Police offender series begins, whereas those born in the first part of 1992 turn 17 before June 2009. At the upper boundary, individuals born in December 1998 reach age 27 in December 2025, so their complete age-17–26 window falls within the available series. Operationally, the full window runs from the month in which an individual turns 17 up to, but excluding, the month in which they turn 27. Restricting the sample to the 1993–1998 cohorts consequently provides six complete birth cohorts, each observed over the same ten-year age window, from the 17th birthday to immediately before the 27th birthday. This design avoids unequal age-based follow-up while retaining multiple cohorts for comparisons across relatively small gender-minority populations. Figure 3 illustrates the relationship between birth cohort and calendar-year coverage. For each offense record, perpetrator's age is calculated from the difference between the earliest recorded occurrence date and the individual's birth date. We retain records with an age of at least 17 and

below 27. Thus, the age restriction is implemented at monthly rather than daily precision.

To limit differences in the length of exposure to New Zealand administrative records and because offenses committed abroad are not observed, we identify extended absences during the outcome window by using the IDI overseas-spell data, which provides the exact date when a person exits and enters New Zealand.[3] The extraction retains overseas spells lasting more than 30 days. These spells are restricted to their overlap with ages 17–26, converted to monthly intervals, and summed within individuals. We exclude individuals who were overseas for 12 months or longer between ages 17–26.[4] This restriction reduces prolonged overseas exposure but does not require continuous presence in New Zealand throughout the observation window.

2.4 Police-recorded offending outcomes

We link the Census population to Police offender records available from June 2009 through December 2025. Using the associated IDI metadata, we map offense codes to the Australian and New Zealand Standard Offence Classification (ANZSOC). We retain assault, harm or endangerment of persons, homicide, robbery, blackmail and extortion, sexual offenses, and theft. The analysis therefore concerns a selected set of interpersonal and victim-related offenses rather than all Police-recorded offending.

Our outcomes measure the extensive margin. For each individual, an indicator equals one if at least one qualifying record is observed during the relevant age window (17–26 in the main analysis) and zero otherwise. Based on the Police offender record entries, we construct the following four main outcomes:

- *Any offense* indicates at least one record in any of the selected offense categories.
- *Assault-related offending* includes assault, harm or endangerment of persons, homicide, and sexual offenses.
- *Serious offending* indicates at least one qualifying record with a seriousness score of at least 100 in the linked offense-code metadata.[5]

[3] A person who has not returned to New Zealand is given the reentry-year of *9999*.

[4] Appendix Table A.2 repeats the analysis using overseas-absence thresholds of 0, 3, 6, 9, and 24 months. The substantive pattern persists: all four gender-minority groups have lower adjusted odds of a qualifying record than cisgender men. Transgender women have a 6.7–8.3 percentage point lower adjusted probability than cisgender men across these specifications (all $p < 0.01$), while their differences from cisgender women remain statistically insignificant. The comparison between transgender men and cisgender men is significant at the 10% level under the three-month threshold and at the 5% level under the other thresholds.

[5] The Ministry of Justice's Justice Sector Seriousness Score measures the relative seriousness of offense types using adult court sentencing data. Scores express average sentence severity in equivalent days of imprisonment, with community-based sentences and fines converted to imprisonment equivalents. Higher scores indicate offenses that typically attract more severe sentences. The score describes the offense type rather than the sentence received by a particular individual. For reference: common manual assault has a seriousness score of 17.83

- *Court action* refers to the method of proceeding recorded by New Zealand Police. Police responses depend on offense characteristics, such as type and seriousness, and offender characteristics, such as age. Three broad response categories are distinguished: *non-court action*, including formal or informal warnings and referrals to a non-court conference, such as a Family Group Conference; *charge and diversion*, through which eligible individuals may resolve charges under the Police Adult Diversion Scheme without receiving a criminal conviction; and *court action*, which involves the Police initiates a prosecution by filing a charging document with a court under the Criminal Procedure Act 2011. We construct a binary indicator equal to one if an individual has at least one qualifying record for which the Police-recorded method of proceeding is court action. This indicator does not establish that the individual was subsequently convicted.

2.5 Victim characteristics and offense circumstances

The Police offender records hold further information on other offense-related characteristics:

- *Offender–victim relationship:* The offender–victim relationship identifies three separate variables, each indicating an offense in which the victim is a family member (including partners, parents and step-parents, children and stepchildren, siblings, boyfriends or girlfriends, grandparents, grandchildren, and relatives providing or receiving care), other known people (including former partners, former boyfriends or girlfriends, flatmates or boarders, friends, non-relative caregivers or care receivers, and other relationships recorded as known but not elsewhere classified), or strangers.

- *Location:* We construct separate indicators for qualifying records in home or residential settings and in selected outside locations. The latter include streets and footpaths, retail and recreational locations, educational and health facilities, and transport-related locations, as identified by Police location codes. Outside locations are not the complement of home or residential locations, and individuals may have records in both categories.

We also link Police offender records to Police victimization records using shared occurrence and offense identifiers. These linkages allow us to construct separate indicators for qualifying records involving victims recorded as female or male. Because the victimization series begins in 2014, these outcomes are measured over the common age window 21–26. The indicators refer to the victim's recorded sex (either female or male) rather than gender identity. They are not mutually exclusive, and a qualifying offender record without a matched victim record of the relevant sex does not contribute to that indicator.

2.6 Family background and residential information

We link each individual to up to two recorded parents using the DIA birth record entries. This means (1) we restrict the sample to those who are born in New Zealand, and (2) we identify recorded parental relationships established at birth rather than the adults with whom the individual lived throughout childhood. This enables us to create consistent measures on income and offense histories of the identified parents.

Parental income is drawn from the total-income module and includes income from wages & salaries, self employment, investments, tax credit, ACC income, student allowances, paid parental leave, income support payments, and other income.[6] For an individual born in calendar year b, parental income is the annual income reported between 1 April of year $b + 14$ and 31 March of year $b + 15$, the boundaries of the financial year. This tax-year-based measure captures parental income two years before the start of the outcome window. Income is adjusted using the consumer price index and summed across available records for the parents identified through the birth registration.[7] Individuals are ranked by combined parental income within year-month-of-birth cohorts and allocated to ten approximately equal-sized groups.

Parental sentencing history is derived from MoJ court-charge records. We retain convictions recorded from the individual's birth up to, but excluding, their 15th birthday for which a recorded parent received a monetary sentence, another community sentence, home detention, or imprisonment. For each recorded parent, following the ordering described in the 2002 Sentencing Act, we retain the highest-ranked sentence within this period. We distinguish individuals with no parent sentenced, one parent sentenced, and both parents sentenced. An additional indicator identifies whether either parent received a sentence in one of the three highest sentencing categories of other community sentence, home detention, or imprisonment.

In addition to the parental resources variables, location of residence is obtained from the IDI address-notification data. The IDI address-notification data combine address records from multiple government agencies, including Inland Revenue and the ministries of Health, Education and Social Development. Records link individuals to de-identified

[6] *ACC income:* Income from ACC (Accident Compensation Corporation) refers to weekly compensation that replaces earnings lost when a covered injury prevents or limits a person's ability to work. These payments are taxable and usually replace up to 80% of pre-injury earnings. They can relate to injuries sustained at work or elsewhere. *Income support payments*: Income support payments from the Ministry of Social Development (MSD) provide financial assistance to help meet living costs. The first pillar is main benefits, which are regular, generally income-tested payments that provide basic income to people who cannot fully support themselves through paid work, including because of unemployment, health limitations or caring responsibilities. Examples include Jobseeker Support, Sole Parent Support and Supported Living Payment. The second pillar is supplementary benefits, which are non-taxable payments that help meet specific ongoing costs, such as housing, disability-related expenses and childcare. Examples include the Accommodation Supplement, Disability Allowance and Childcare Subsidy.

[7] The share of parents with no income recorded is <1% and we set their income to 0. As we include income from self-employment and investment, the total income can also be negative.

address identifiers and notification dates, enabling researchers to construct approximate residential histories. We identify the postcode associated with the address record active 30 days before the individual's 15th birthday. Standard errors are clustered at this postcode level.

The final analysis dataset comprises Census self-respondents who satisfy the birth-linkage, residential-information, and overseas-exposure restrictions reported in Table A.1. The adjustment variables comprise year-month of birth, Census ethnicity, parental-income group, the number of recorded parents sentenced, and an indicator for a higher-category parental sentence.

Table A.1 documents the sequential construction of the analysis sample. The initial population comprises 418,785 individuals from the 1993–1998 birth cohorts who were resident in New Zealand in March 2023. Restricting the sample to New Zealand-born individuals linked to DIA birth registrations retains 257,460 individuals; requiring Census self-responses on both sex at birth and gender reduces this to 206,682. Residential postcode information is identified for 199,200 individuals. Excluding those with 12 or more recorded months overseas at ages 17–26 yields the final sample of 170,775, or 40.8% of the initial population. The estimates therefore describe New Zealand-born Census self-respondents meeting the linkage and residence restrictions.

2.7 Descriptive statistics

Table 1 presents sample characteristics and unadjusted offending outcomes by gender identity. Cisgender men and women account for 48.70% and 49.75% of the sample, respectively. The remaining groups comprise 318 transgender men (0.19%), 429 transgender women (0.25%), 762 non-binary people assigned male at birth (AMAB; 0.45%), and 1,140 non-binary people assigned female at birth (AFAB; 0.67%).

Background characteristics differ across the groups. The proportion identifying as European only ranges from 59.75% among cisgender women to 69.53% among non-binary people AFAB. Transgender men are overrepresented in the lowest two parental-income groups: 25.16% of transgender men fall into these groups, compared with 19.63% of cisgender men and 20.52% of cisgender women. They also have the lowest representation in the highest two income groups, at 15.41%. Parental sentencing histories show further variation. Both recorded parents had received a sentence for 6.92% of transgender men and 6.76% of transgender women, compared with 4.80% of cisgender men and 5.69% of cisgender women. Non-binary people AFAB have the highest proportion with neither recorded parent sentenced (80.51%) and the lowest proportion with a parent receiving a higher-category sentence (10.45%).

Overall, 8.96% of individuals have at least one qualifying Police-recorded offense at ages 17–26. The proportion is highest among cisgender men (11.81%), followed by transgender

men (8.49%), cisgender women (6.29%), non-binary people AMAB (6.03%), transgender women (5.36%), and non-binary people AFAB (3.60%). Thus, all four gender-minority groups have lower unadjusted proportions than cisgender men. Transgender women and non-binary people AMAB have proportions closer to that of cisgender women, while non-binary people AFAB have the lowest proportion.

These patterns are consistently observed for different age ranges (in which the age range is cut in half to measure any qualifying record at ages 17–21 and 22–26 separately), for serious offenses, and regardless of victim's characteristics or location of the crime. In all cases, cisgender men consistently have the highest unadjusted proportions. The following regression analysis examines whether these differences persist after adjustment for ethnicity, family background, and year-month of birth.

3 Empirical approach

3.1 Regression specification

We estimate a separate logistic regression for each binary outcome. For individual i and outcome k, let Y_{ik} equal one if at least one qualifying record is observed within the relevant window and zero otherwise. The main window covers ages 17–26. Age-specific analyses use Y_{ik} measured within ages 17–21 and 22–26 separately, while outcomes involving victim sex are only measured for ages 21–26. The model is

$$\Pr(Y_{ik} = 1 \mid G_i, \mathbf{X}_i, b_i) = \Lambda\left(\alpha_k + \sum_{g \neq \mathrm{CM}} \beta_{kg} \mathbf{1}\{G_i = g\} + \mathbf{X}_i' \boldsymbol{\gamma}_k + \delta_{k,b_i}\right), \tag{1}$$

where $\Lambda(z) = \exp(z)/(1 + \exp(z))$ is the logistic distribution function. G_i denotes the six-category gender identity classification, with cisgender men (CM) as the reference group. The remaining categories are cisgender women, transgender men, transgender women, and non-binary people assigned male or female at birth.

The vector $\mathbf{X}_i$ contains categorical indicators for Census ethnicity, parental-income group, and the number of recorded parents sentenced during the individual's first 15 years of life, together with an indicator for whether either parent received a sentence in one of the three highest sentence categories. Parental income is grouped into ten approximately equal-sized categories within year-month of birth, as described in Section 2. The terms δ_{k,b_i} are fixed effects for year-month of birth, allowing baseline outcome probabilities to vary across the 72 birth-month cohorts included in the sample.

Within each outcome specification, individuals share the same nominal age window, although the calendar periods covered differ by birth cohort and permitted overseas absences reduce actual exposure to New Zealand administrative recording. The year-month-of-birth fixed effects account for differences in baseline risk across these cohorts

without imposing a linear cohort trend. Standard errors are clustered by residential postcode measured 30 days before the individual's 15th birthday. This permits within-postcode dependence in the error term.

3.2 Adjusted probabilities and probability differences

Our principal results are adjusted predicted probabilities, which express group differences on the probability scale. For outcome k and gender-identity group g, we calculate

$$\widehat{p}_{kg} = \frac{1}{N_k} \sum_{i \in \mathcal{S}_k} \Lambda \left(\widehat{\alpha}_k + \widehat{\beta}_{kg} + \mathbf{X}_i' \widehat{\boldsymbol{\gamma}}_k + \widehat{\delta}_{k,b_i} \right), \tag{2}$$

where $\mathcal{S}_k$ is the estimation sample, N_k is its size, and $\widehat{\beta}_{k,\mathrm{CM}} = 0$. For each gender-identity category, we calculate fitted probabilities for every individual in the estimation sample, retaining that individual's observed covariates and birth cohort, and average these probabilities. Each group's prediction is therefore standardized to the same covariate distribution. These estimates summarize adjusted associations and differ from both unadjusted group proportions and predictions evaluated at mean covariate values.

Additionally, we directly compare transgender women (TF) with cisgender men and cisgender women (CF):

$$\widehat{\Delta}_{k,\mathrm{TF-CM}} = \widehat{p}_{k,\mathrm{TF}} - \widehat{p}_{k,\mathrm{CM}}, \tag{3}$$

$$\widehat{\Delta}_{k,\mathrm{TF-CF}} = \widehat{p}_{k,\mathrm{TF}} - \widehat{p}_{k,\mathrm{CF}}. \tag{4}$$

Multiplying these differences by 100 expresses them in percentage points. Standard errors for adjusted probabilities and their differences are obtained using the delta method and the postcode-clustered covariance matrix. Tests of probability differences account for the covariance between the two predictions. The relevant null hypothesis for a group comparison is $H_0 : \Delta_{k,g-h} = 0$; a test that an individual group's probability equals zero does not test a difference between groups.

3.3 Adjusted odds ratios and inference

We also report adjusted odds ratios to summarize the conditional associations from the logistic regressions. For group g relative to cisgender men, the adjusted odds ratio is

$$\widehat{\mathrm{aOR}}_{kg} = \exp(\widehat{\beta}_{kg}). \tag{5}$$

Values below one indicate lower conditional odds of a qualifying record than among cisgender men, while values above one indicate higher conditional odds. The comparison

between transgender women and cisgender women is calculated directly as

$$\widehat{\mathrm{aOR}}_{k,\mathrm{TF:CF}} = \exp\left(\widehat{\beta}_{k,\mathrm{TF}} - \widehat{\beta}_{k,\mathrm{CF}}\right). \tag{6}$$

We test equality of their conditional odds using the Wald test of $H_0 : \beta_{k,\mathrm{TF}} - \beta_{k,\mathrm{CF}} = 0$. Reported odds-ratio standard errors use the delta method, whereas significance tests are conducted on the underlying log-odds scale. All contrasts are calculated from unrounded estimates. The adjusted probabilities and odds ratios provide complementary summaries. The former express standardised outcome probabilities and absolute differences, whereas the latter describe conditional differences in odds.

4 Results

4.1 Any recorded offense and age-specific outcomes

Table 2 examines the sensitivity of the estimated gender differences to sequential covariate adjustment. Model I includes gender-identity indicators only. Models II–V successively add year-month-of-birth fixed effects, ethnicity, parental-income group, and parental sentencing measures. All specifications use the same 170,775 individuals, allowing comparisons across models without changes in sample composition.

The broad pattern is stable across specifications. Transgender women's predicted probability of a qualifying record changes from 5.4% in Model I to 5.1% in Model V, while the corresponding estimate for cisgender men increases from 11.8% to 12.2%. Their estimated probability difference changes from −6.5 to −7.2 percentage points and is statistically significant at the 1% level in every model. Differences between transgender women and cisgender women remain small, ranging from −1.0 to −0.7 percentage points, and none is statistically significant. The addition of the observed covariates therefore preserves the main comparison between transgender women and the two cisgender groups.

For transgender men, the odds ratio relative to cisgender men declines from 0.695 in Model I to 0.617 in Model V. This comparison is significant at the 10% level in Models I–III and at the 5% level in Models IV–V, following the inclusion of parental income. For non-binary people AFAB, adjustment raises the predicted probability from 3.6% to 4.6%, although their odds remain significantly below those of cisgender men in every specification. Estimates for non-binary people AMAB change comparatively little. Changes across specifications describe sensitivity to adjustment rather than the proportion of a gender difference explained by each covariate block. In particular, odds ratios from nested logistic models are not a simple decomposition of the unadjusted association.

The age-specific estimates in Table 3 show a similar pattern. Transgender women's adjusted probabilities are 4.2% at ages 17–21 and 2.4% at ages 22–26, compared with

9.4% and 5.8% among cisgender men. The corresponding differences are −5.2 and −3.4 percentage points, respectively, both significant at the 1% level. Differences from cisgender women are −0.6 and 0.1 percentage points, neither statistically significant. The comparison between transgender women and cisgender men therefore persists in both age windows. Both subwindows cover five years, so their identical lengths do not explain the lower probabilities in the later window. Nevertheless, these probabilities are not annual offending rates, and the analysis does not isolate age effects from calendar-period influences.

4.2 Offense type and court action

Table 4 distinguishes assault-related offending, serious offending, and court action. The patterns observed for any selected offenses in earlier tables are similarly observed for all outcomes. Gender minorities are substantially less likely to have assault-related offenses, have serious offenses, or have court proceedings for their offenses. The adjusted probabilities among cisgender men are 9.6%, 6.9%, and 9.7%, respectively. Among transgender women, they are 4.1%, 2.6%, and 3.2%. The corresponding differences are −5.5, −4.3, and −6.4 percentage points, all significant at the 1% level. Adjusted odds ratios for transgender women relative to cisgender men range from 0.284 for court action to 0.381 for assault-related offending.

The differences between transgender women and cisgender women are small in absolute terms: 0.2 percentage points for assault-related offending, 0.8 for serious offending, and −0.5 for court action. None is statistically significant. The odds-ratio comparisons likewise do not reject equality. All other gender-minority groups also have point estimates below those of cisgender men for these outcomes, although the precision varies; for example, the assault-related odds ratio for transgender men is significant only at the 10% level. Court action denotes a Police-recorded proceeding involving court action and does not establish a subsequent conviction.

4.3 Victim characteristics and offense location

Table 5 presents qualifying records at ages 21–26 involving victims recorded as female or male. The adjusted probabilities among transgender women are 1.5% and 1.3%, respectively, compared with 3.4% and 3.3% among cisgender men. The corresponding differences are −1.9 and −2.0 percentage points, both significant at the 1% level. Differences from cisgender women are 0.4 and 0.1 percentage points, respectively, and neither is statistically significant. These outcomes require matched victimization records and use a shorter age window than the main outcomes.

Table 6 distinguishes family members, other known people, and strangers. Again, we observe similar patterns in which cisgender men have the highest probabilities to have offenses regardless of victim's relationship to the perpetrator. We do not observe any

relationship between perpetrator's gender and victim's relationship to the perpetrator. For transgender women, the adjusted probabilities are 2.2%, 1.9%, and 1.7%, respectively. The corresponding differences from cisgender men are −1.5, −2.1, and −2.6 percentage points. Each probability difference is significant at the 5% level, although the conditional odds ratio comparison for family-related offending is significant only at the 10% level. None of the probability differences or odds-ratio comparisons between transgender women and cisgender women is statistically significant.

Finally, Table 7 distinguishes home or residential locations from selected outside locations. The patterns observed in other analyses that cisgender men are most likely to hold offending records and the lack of systemic link between gender minorities and characteristics of the crime again hold. Transgender women's adjusted probabilities are 2.8% and 3.5%, compared with 6.8% and 8.3% among cisgender men. The differences of −4.0 and −4.8 percentage points are significant at the 1% level. Relative to cisgender women, the estimated differences round to 0.0 and −0.7 percentage points and are not statistically significant.

Across the reported outcomes, transgender women have lower adjusted probabilities than cisgender men, while the comparisons with cisgender women do not reject equality at the 5% level. The disaggregated results describe recorded outcomes in different circumstances; they do not establish that groups are equivalent or identify the mechanisms producing the observed differences.

4.4 Gender-minority people AMAB with hormone prescriptions

As a supplementary analysis, we examine transgender women and non-binary people assigned male at birth (AMAB) with at least one recorded dispensing of estrogen and a testosterone-blocking medication at ages 17–26. We identify these dispensings in the Pharmaceutical Collection. The subgroup contains 342 individuals, whom we compare with cisgender men and cisgender women in an estimation sample of 168,468 individuals. The dispensing criterion provides additional administrative information but does not by itself establish treatment indication or use.

We re-estimate the model for any qualifying Police-recorded offense at ages 17–26 using the main adjustment variables and postcode-clustered standard errors. Adjusted probabilities are averaged over this supplementary estimation sample, so comparisons with the main results reflect both subgroup selection and a different standardization population.

Table A.3 reports an adjusted probability of 5.4% for the hormone-dispensing subgroup, compared with 12.2% for cisgender men and 6.1% for cisgender women. The corresponding differences are −6.9 percentage points ($p < 0.01$) and −0.7 percentage points (not statistically significant). Adjusted odds ratios are 0.375 relative to cisgender men and

0.871 relative to cisgender women.

The broad pattern resembles the main findings, but this analysis combines transgender women and non-binary people AMAB and does not isolate differences within either group. Dispensing and offending are measured within the same age window, without establishing their temporal ordering. The estimates therefore do not identify effects of hormone treatment or other forms of gender affirmation.

4.5 School performance and police-recorded offending

We examine whether gender differences in Police-recorded offending are also evident within groups defined by early school performance. We use the Ministry of Education's Student Expected Percentile Score (reported here on a 0–1 scale), also termed prior education performance, which summarizes students' National Certificate of Educational Achievement (NCEA; New Zealand's main official metric for measuring student achievement) results relative to other students' performance (Ussher, 2008, Appendix C). We select individuals whose first recorded Level 1 score is observed at ages 14–17 and we match the score records for 145,272 individuals (85.1%), with 2,751 (1.9%) having their score at age 14, 44,868 (30.9%) at age 15, 82,857 (57.0%) at a age 16, and 14,796 (10.2%) at age 17.

We distinguish a low score (belonging to the bottom 3 deciles) from a higher score (belonging to the top 7 deciles). We exclude scores coded -1, indicating that no Level 1 standards have been taken, and retain scores coded zero, indicating that only not-achieved results are recorded. The resulting sample contains 145,272 individuals, compared with 170,775 in the main analysis. We retain the main adjustment variables and postcode-clustered standard errors.

Table A.4 shows that cisgender women, transgender men, and both non-binary groups have higher unadjusted mean educational performance scores than cisgender men, with all four differences statistically significant ($p \leq 0.012$). Non-binary people assigned female at birth have the highest mean score (0.488), compared with 0.364 among cisgender men. Transgender women's mean score is 0.370, and the difference relative to cisgender men is not statistically significant ($p = 0.541$).

Table A.5 first reports results for the pooled available-score sample. All four gender-minority groups have significantly lower adjusted odds of a qualifying record than cisgender men, with adjusted odds ratios between 0.338 and 0.484. Transgender women's adjusted probability is 4.6%, compared with 10.3% among cisgender men and 5.0% among cisgender women. Their estimated differences are -5.8 percentage points relative to cisgender men ($p < 0.01$) and -0.4 percentage points relative to cisgender women (not statistically significant).

In the low score group, transgender women have an adjusted probability 11.8 percentage points below cisgender men ($p < 0.01$) and 4.1 percentage points below cisgender women

($p < 0.05$), although the latter odds-ratio comparison is not statistically significant. In the higher score group, the corresponding differences are -1.5 and 1.3 percentage points, neither statistically significant. Both non-binary groups have significantly lower adjusted odds than cisgender men in each score category; estimates for transgender men in the higher category are not reported due to small count.

5 Conclusion

This study documents substantial variation in Police-recorded offending across groups defined jointly by sex at birth and gender recorded in the 2023 Census. Over ages 17–26, all four gender-minority groups have lower adjusted odds of a qualifying record than cisgender men. Transgender women also have lower adjusted probabilities than cisgender men across the main outcomes distinguished by offense type, court action, victim characteristics, and location. Corresponding comparisons with cisgender women do not reject equality at the 5% level. These results do not support a general characterization of transgender women as having higher recorded offending in the outcomes examined.

The supplementary analyses qualify this overall pattern. Among individuals in the lower school-performance category, transgender women have a significantly lower adjusted probability than cisgender women, although the corresponding odds-ratio test does not reject equality. In the higher school-performance category, neither probability difference from the cisgender comparison groups is statistically significant. Estimates for the hormone-dispensing subgroup and alternative overseas-absence thresholds broadly resemble the main findings, but refer to differently selected samples.

Several limitations affect interpretation. First, Police records capture offenses that enter the recording system and are attributed to an identified person. Group differences may reflect reporting, detection, and attribution as well as underlying behavior. The outcomes cover selected offense categories and measure whether a person has any qualifying record; they do not measure total offending frequency, cumulative harm, or subsequent convictions.

Second, gender is measured in the 2023 Census, whereas some qualifying records precede that measurement. The classification does not establish an individual's gender identity at the time of an earlier offense or the timing of gender affirmation. The analysis therefore relates Census-recorded gender to administrative offending histories.

Third, adjustment for observed demographic and family-background characteristics does not eliminate unobserved differences in individual circumstances, offending opportunities, or contact with the criminal justice system. The estimates are descriptive associations rather than causal effects of gender identity, gender affirmation, or hormone treatment. Small group sizes also limit precision, particularly for less frequent outcomes. Failure to reject a difference should not be interpreted as evidence of equivalence.

Fourth, the estimates apply to New Zealand-born Census self-respondents meeting the linkage and residence restrictions. They need not represent people excluded by these requirements. The overseas restriction limits prolonged absences but does not ensure continuous presence in New Zealand. Outcomes involving victim sex additionally require linked victimization records and are measured at ages 21–26.

Finally, the analysis concerns the 1993–1998 birth cohorts during late adolescence and early adulthood, with main outcome windows falling within 2010–2025. The results do not establish patterns for older ages, earlier cohorts, or other institutional settings. Replication using comparable measures of sex at birth, gender, and recorded offending would help establish how the findings vary across populations and criminal justice systems.

Data deposition information

The individual-level data used in this study are held in the Integrated Data Infrastructure (IDI), maintained by Statistics New Zealand (Stats NZ). These data are subject to statutory confidentiality restrictions, including those under the Data and Statistics Act 2022, and cannot be deposited in a public repository or shared directly by the authors. Access is available only to approved researchers working on approved projects within a secure Stats NZ Data Lab environment. Researchers wishing to access the underlying data must apply directly to Stats NZ and satisfy its access requirements. Information about the application process is available at `https://www.stats.govt.nz/integrated-data/apply-to-use-microdata-for-research/`. All research outputs are subject to confidentiality checking before release.

References

Adams, A., Huttunen, K., Nix, E. and Zhang, N. (2024). The dynamics of abusive relationships, *The Quarterly Journal of Economics* **139**(4): 2135–2180.

Adams, A., Huttunen, K., Nix, E. and Zhang, N. (forthcoming). The economic impacts of sexual assault, *American Economic Review* .

Adams-Prassl, A., Huttunen, K., Nix, E. and Zhang, N. (2024). Violence against women at work, *The Quarterly Journal of Economics* **139**(2): 937–991.

Adriaenssen, A., Paoli, L., Karstedt, S., Visschers, J., Greenfield, V. A. and Pleysier, S. (2020). Public perceptions of the seriousness of crime: Weighing the harm and the wrong, *European Journal of Criminology* **17**(2): 127–150.
URL: *https://doi.org/10.1177/1477370818772768*

Bell, B., Bindler, A. and Machin, S. (2018). Crime scars: Recessions and the making of career criminals, *Review of Economics and Statistics* **100**(3): 392–404.

Berlingieri, F. and Kovacic, M. (2025). Health and relationship quality of sexual minorities in europe, *Journal of Population Economics* **38**: 15. Article 15.

Campaniello, N. and Gavrilova, E. (2018). Uncovering the gender participation gap in crime, *European Economic Review* **109**: 289–304.

Carpenter, C. S., Eppink, S. T. and Gonzales, G. (2020). Transgender status, gender identity, and socioeconomic outcomes in the united states, *ILR Review* **73**(3): 573–599.

Carpenter, C. S., Goodman, L. and Lee, M. J. (2024). Transgender earnings gaps in the united states: evidence from administrative data, *Technical report*, National Bureau of Economic Research.

Carpenter, C. S., Kirkpatrick, L., Lee, M. J. and Plum, A. (2025). Economic outcomes of gender diverse people: New evidence from linked administrative data in new zealand, *Economics Letters* **247**: 112155.

Carpenter, C. S., Lee, M. J. and Nettuno, L. (2022). Economic outcomes for transgender people and other gender minorities in the united states: First estimates from a nationally representative sample, *Southern Economic Journal* **89**(2): 280–304.

Cornaglia, F., Feldman, N. E. and Leigh, A. (2014). Crime and mental well-being, *Journal of Human Resources* **49**(1): 110–140.

Currie, J. and Tekin, E. (2012). Understanding the cycle: Childhood maltreatment and future crime, *Journal of Human Resources* **47**(2): 509–549.

Damm, A. P. and Dustmann, C. (2014). Does growing up in a high crime neighborhood affect youth criminal behavior?, *American Economic Review* **104**(6): 1806–1832.

Dasgupta, K., Diegmann, A., Kirchmaier, T. and Plum, A. (2022). The gender reveal: The effect of sons on young fathers' criminal behavior and labor market activities, *Labour Economics* **78**: 102224.

Deming, D. J. (2011). Better schools, less crime?, *Quarterly Journal of Economics* **126**(4): 2063–2115.

Dhejne, C., Lichtenstein, P., Boman, M., Johansson, A. L., Långström, N. and Landén, M. (2011). Long-term follow-up of transsexual persons undergoing sex reassignment surgery: cohort study in sweden, *PloS one* **6**(2): e16885.

Flores, A. R., Meyer, I. H., Langton, L. and Herman, J. L. (2021). Gender identity disparities in criminal victimization: National Crime Victimization Survey, 2017–2018, *American Journal of Public Health* **111**(4): 726–729.

Heckman, J., Pinto, R. and Savelyev, P. (2013). Understanding the mechanisms through which an influential early childhood program boosted adult outcomes, *American Economic Review* **103**(6): 2052–2086.

Hederos Eriksson, K., Hjalmarsson, R., Lindquist, M. J. and Sandberg, A. (2016). The importance of family background and neighborhood effects as determinants of crime, *Journal of Population Economics* **29**: 219–262.

Hjalmarsson, R. and Lindquist, M. J. (2013). The origins of intergenerational associations in crime: Lessons from Swedish adoption data, *Labour Economics* **20**: 68–81.

Saxby, K., Hutchinson Tovar, S., Bishop, G. M., Down, I., Spencer, R., Petrie, D. and Aitken, Z. (2025). Gender identity and mental health inequalities 2001–2022: Population-level evidence from an Australian cohort study, *BMJ Mental Health* **28**(1): e301277.

Treshenkov, V. T., Turjeman, H. and Dagan, N. (2026). Measuring the public's perceptions of crime seriousness: A review and the development of a scale, *Journal of Criminology* **59**(2): 204–230.
URL: *https://doi.org/10.1177/26338076251364561*

Twenge, J. M., Wells, B. E., Le, J. and Rider, G. N. (2025). Increases in self-identifying as transgender among US adults, 2014–2022, *Sexuality Research and Social Policy* **22**(2): 755–773.

Ussher, S. (2008). Post-school choices: How well does academic achievement predict the tertiary education choices of school leavers?, *Technical report*, Ministry of Education.
URL: *https://www.educationcounts.govt.nz/publications/tertiary*$_e$*ducation/education* − *learners*/23103

Table 1: Sample characteristics and unadjusted Police-recorded offending by gender identity

	All	Cisgender men	Cisgender women	Transgender men	Transgender women	Non-binary AMAB	Non-binary AFAB
Sample composition							
Individuals, N	170,775	83,172	84,954	318	429	762	1,140
Share of total sample (%)	100.0	48.70	49.75	0.19	0.25	0.45	0.67
Background characteristics							
European only	61.57	63.24	59.75	66.67	64.80	67.23	69.53
Parental income: lowest two groups	20.08	19.63	20.52	25.16	18.41	20.97	18.88
Parental income: middle six groups	60.78	60.47	61.17	59.43	60.37	57.01	57.59
Parental income: highest two groups	19.14	19.90	18.31	15.41	21.21	22.02	23.53
Neither recorded parent sentenced	72.75	74.00	71.42	70.44	70.86	74.18	80.51
One recorded parent sentenced	22.02	21.20	22.89	22.64	22.38	20.71	17.03
Both recorded parents sentenced	5.23	4.80	5.69	6.92	6.76	5.11	2.46
Either parent: higher-category sentence	14.73	13.81	15.68	16.04	16.78	14.42	10.45
Police-recorded offending (%)							
Any offense	8.96	11.81	6.29	8.49	5.36	6.03	3.60
Any offence, ages 17–21	6.93	9.05	4.95	7.55	4.90	4.33	2.63
Any offence, ages 22–26	3.91	5.52	2.39	2.83	2.10	3.15	1.32
Assault-related offence	6.66	9.32	4.14	6.60	4.90	4.72	2.37
Serious offence	4.19	6.63	1.88	1.89	2.80	3.15	0.79
Court action	6.53	9.33	3.88	5.66	3.50	4.72	2.37
Female victim	2.17	3.24	1.14	2.20	1.63	1.57	0.88
Male victim	2.18	3.17	1.24	2.20	1.40	2.23	0.44
Victim: family member	2.48	3.48	1.55	1.89	2.10	1.18	0.53
Victim: other known person	2.74	3.86	1.67	1.89	2.10	2.76	0.79
Victim: stranger	2.74	4.11	1.44	1.89	1.40	1.97	1.05
Location: home or residential	4.69	6.55	2.92	5.66	2.80	3.15	1.58
Location: selected outside locations	6.13	8.03	4.35	4.72	3.50	4.72	2.37

Notes: Sample: birth cohorts 1993–1998. AMAB and AFAB denote assigned male and female at birth. Parental-income groups are constructed within year-month of birth; the ten groups are combined here for presentation. Parental sentencing is measured from the individual's birth through age 15; higher-category sentences comprise the three highest categories: other community, home-detention, and imprisonment. Offending outcomes indicate at least one qualifying record at ages 17–26 unless otherwise stated, using all individuals in the column as the denominator. Variable definitions and sample restrictions are described in Section 2.

Table 2: Gender identity and police-recorded offending: sequential covariate adjustment

	Any selected police-recorded offense				
Model	I	II	III	IV	V
Panel A: (Adjusted) predicted probabilities					
Cisgender men	0.118***	0.118***	0.121***	0.121***	0.122***
	(0.003)	(0.003)	(0.002)	(0.002)	(0.001)
Cisgender women	0.063***	0.063***	0.061***	0.061***	0.060***
	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)
Transgender men	0.085***	0.085***	0.087***	0.083***	0.082***
	(0.016)	(0.016)	(0.016)	(0.015)	(0.015)
Transgender women	0.054***	0.055***	0.053***	0.054***	0.051***
	(0.011)	(0.011)	(0.010)	(0.010)	(0.010)
non-binary, assigned male at birth	0.060***	0.061***	0.065***	0.065***	0.064***
	(0.009)	(0.009)	(0.010)	(0.009)	(0.009)
non-binary, assigned female at birth	0.036***	0.037***	0.042***	0.043***	0.046***
	(0.006)	(0.006)	(0.007)	(0.007)	(0.007)
Differences in (adjusted) predicted probabilities					
Transgender women − cisgender men	−0.065***	−0.063***	−0.067***	−0.067***	−0.072***
	(0.010)	(0.011)	(0.010)	(0.010)	(0.010)
Transgender women − cisgender women	-0.009	-0.008	-0.008	-0.007	−0.010
	(0.010)	(0.011)	(0.010)	(0.010)	(0.009)
Panel B: (Adjusted) odds ratios					
Cisgender women	0.501***	0.498***	0.461***	0.451***	0.431***
	(0.010)	(0.010)	(0.009)	(0.009)	(0.009)
Transgender men	0.695*	0.692*	0.685*	0.641**	0.617**
	(0.140)	(0.141)	(0.143)	(0.136)	(0.132)
Transgender women	0.423***	0.432***	0.396***	0.394***	0.351***
	(0.088)	(0.090)	(0.083)	(0.082)	(0.074)
non-binary, assigned male at birth	0.479***	0.484***	0.496***	0.486***	0.463***
	(0.075)	(0.076)	(0.080)	(0.078)	(0.075)
non-binary, assigned female at birth	0.279***	0.285***	0.304***	0.306***	0.316***
	(0.048)	(0.049)	(0.054)	(0.054)	(0.057)
Transgender women versus cisgender women	0.844	0.867	0.859	0.873	0.816
	(0.175)	(0.179)	(0.179)	(0.181)	(0.171)
Panel C: Added covariates					
Year-month-of-birth FE	✗	✓	✓	✓	✓
Ethnicity	✗	✗	✓	✓	✓
Parental income	✗	✗	✗	✓	✓
Parental sentencing	✗	✗	✗	✗	✓
Observations	170,775	170,775	170,775	170,775	170,775

Notes: All models use the same sample of 170,775 New Zealand-born individuals from the 1993–1998 birth cohorts. The outcome indicates any qualifying Police record at ages 17–26. Model I includes gender-identity indicators only. Model II adds year-month-of-birth fixed effects; Model III additionally includes Census ethnicity; Model IV adds parental-income group; and Model V adds parental sentencing measures. Parental sentencing comprises the number of recorded parents sentenced before the individual's 15th birthday and an indicator for either parent receiving a higher-category sentence. Panel A reports predicted probabilities and differences on the 0–1 scale; differences multiplied by 100 are percentage points. Estimates are unadjusted in Model I; predictions in Models II–V are averaged over the common sample's observed covariate distribution. Panel B reports odds ratios relative to cisgender men except in the final comparison; these are unadjusted in Model I and adjusted in Models II–V. Panel C summarizes covariate inclusion. Standard errors clustered by residential postcode measured 30 days before the individual's 15th birthday are in parentheses. Standard errors for probabilities, contrasts, and odds ratios use the delta method. Contrasts use unrounded estimates. Stars test probabilities and differences against zero and odds ratios against one: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 3: Gender identity and police-recorded offending by age window

	Any selected police-recorded offense		
Gender identity	Ages 17–26	Ages 17–21	Ages 22–26
Panel A: Adjusted predicted probabilities			
Cisgender men	0.122***	0.094***	0.058***
	(0.001)	(0.001)	(0.001)
Cisgender women	0.060***	0.048***	0.023***
	(0.001)	(0.001)	(0.001)
Transgender men	0.082***	0.070***	0.034***
	(0.015)	(0.013)	(0.011)
Transgender women	0.051***	0.042***	0.024***
	(0.010)	(0.009)	(0.007)
non-binary, assigned male at birth	0.064***	0.049***	0.034***
	(0.009)	(0.008)	(0.007)
non-binary, assigned female at birth	0.046***	0.034***	0.017***
	(0.007)	(0.006)	(0.005)
Differences in adjusted predicted probabilities			
Transgender women − cisgender men	−0.072***	−0.052***	−0.034***
	(0.010)	(0.009)	(0.007)
Transgender women − cisgender women	−0.010	−0.006	0.001
	(0.009)	(0.009)	(0.007)
Panel B: Adjusted odds ratios			
Cisgender women	0.431***	0.458***	0.361***
	(0.009)	(0.010)	(0.011)
Transgender men	0.617**	0.705	0.556*
	(0.132)	(0.157)	(0.190)
Transgender women	0.351***	0.397***	0.378***
	(0.074)	(0.091)	(0.115)
non-binary, assigned male at birth	0.463***	0.473***	0.560***
	(0.075)	(0.088)	(0.115)
non-binary, assigned female at birth	0.316***	0.313***	0.258***
	(0.057)	(0.061)	(0.074)
Transgender women versus cisgender women	0.816	0.866	1.049
	(0.171)	(0.197)	(0.314)
Observations	170,775	170,775	170,775

Notes: Sample: birth cohorts 1993–1998. Outcomes indicate any qualifying Police record in the stated age window. Ages 17–26 cover ten years, while ages 17–21 and 22–26 each cover five years. Subwindow probabilities are neither annual risks nor additive, because individuals may have qualifying records in both subwindows. Panel A reports average adjusted probabilities and differences on the 0–1 scale; differences multiplied by 100 are percentage points. Panel B reports adjusted odds ratios, relative to cisgender men except in the final comparison. Models adjust for year-month of birth, ethnicity, parental-income group, and parental sentencing measures. Postcode-clustered standard errors are in parentheses; aOR standard errors use the delta method. Contrasts use unrounded estimates. Stars test probabilities and differences against zero and aORs against one: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 4: Gender identity and selected police-recorded offending outcomes

	Police-recorded outcome at ages 17–26		
Gender identity	Assault-related	Serious offense	Court action
Panel A: Adjusted predicted probabilities			
Cisgender men	0.096***	0.069***	0.097***
	(0.001)	(0.001)	(0.001)
Cisgender women	0.040***	0.018***	0.037***
	(0.001)	(0.001)	(0.001)
Transgender men	0.065***	0.022***	0.061***
	(0.014)	(0.007)	(0.013)
Transgender women	0.041***	0.026***	0.032***
	(0.009)	(0.007)	(0.007)
non-binary, assigned male at birth	0.052***	0.034***	0.051***
	(0.008)	(0.007)	(0.008)
non-binary, assigned female at birth	0.030***	0.011***	0.032***
	(0.006)	(0.003)	(0.007)
Differences in adjusted predicted probabilities			
Transgender women − cisgender men	−0.055***	−0.043***	−0.064***
	(0.009)	(0.007)	(0.007)
Transgender women − cisgender women	0.002	0.008	−0.005
	(0.009)	(0.007)	(0.007)
Panel B: Adjusted odds ratios			
Cisgender women	0.365***	0.225***	0.330***
	(0.008)	(0.008)	(0.008)
Transgender men	0.625*	0.273***	0.576**
	(0.151)	(0.100)	(0.143)
Transgender women	0.381***	0.330***	0.284***
	(0.087)	(0.093)	(0.071)
non-binary, assigned male at birth	0.486***	0.450***	0.467***
	(0.085)	(0.093)	(0.082)
non-binary, assigned female at birth	0.269***	0.133***	0.283***
	(0.059)	(0.043)	(0.064)
Transgender women versus cisgender women	1.045	1.469	0.861
	(0.238)	(0.412)	(0.211)
Observations	170,775	170,775	170,775

Notes: Sample: birth cohorts 1993–1998; outcomes at ages 17–26. Assault-related offending includes assault, harm or endangerment, homicide, and sexual offenses. Serious offending uses a seriousness score of at least 100. Court action is a Police proceeding, not a conviction. Panel A reports average adjusted probabilities and differences on the 0–1 scale; differences multiplied by 100 are percentage points. Panel B reports adjusted odds ratios, relative to cisgender men except in the final comparison. Models adjust for year-month of birth, ethnicity, parental-income group, and parental sentencing measures. Postcode-clustered standard errors are in parentheses; aOR standard errors use the delta method. Contrasts use unrounded estimates. Stars test probabilities and differences against zero and aORs against one: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 5: Gender identity and police-recorded offending by victim sex

	Any selected police-recorded offense	
Gender identity	Female victim	Male victim
Panel A: Adjusted predicted probabilities		
Cisgender men	0.034***	0.033***
	(0.001)	(0.001)
Cisgender women	0.011***	0.012***
	(0.000)	(0.000)
Transgender men	0.022***	0.021***
	(0.008)	(0.009)
Transgender women	0.015***	0.013***
	(0.005)	(0.005)
non-binary, assigned male at birth	0.017***	0.025***
	(0.005)	(0.006)
non-binary, assigned female at birth	0.012***	0.006**
	(0.004)	(0.003)
Differences in adjusted predicted probabilities		
Transgender women − cisgender men	−0.019***	−0.020***
	(0.005)	(0.005)
Transgender women − cisgender women	0.004	0.001
	(0.005)	(0.005)
Panel B: Adjusted odds ratios		
Cisgender women	0.298***	0.335***
	(0.013)	(0.012)
Transgender men	0.629	0.627
	(0.240)	(0.284)
Transgender women	0.413**	0.362***
	(0.151)	(0.138)
non-binary, assigned male at birth	0.480**	0.729
	(0.140)	(0.185)
non-binary, assigned female at birth	0.335***	0.168***
	(0.112)	(0.075)
Transgender women versus cisgender women	1.388	1.081
	(0.501)	(0.407)
Observations	170,775	170,775

Notes: Sample: birth cohorts 1993–1998. Outcomes indicate a qualifying record at ages 21–26 involving the stated victim sex. Categories may overlap. Panel A reports average adjusted probabilities and differences on the 0–1 scale; differences multiplied by 100 are percentage points. Panel B reports adjusted odds ratios, relative to cisgender men except in the final comparison. Models adjust for year-month of birth, ethnicity, parental-income group, and parental sentencing measures. Postcode-clustered standard errors are in parentheses; aOR standard errors use the delta method. Contrasts use unrounded estimates. Stars test probabilities and differences against zero and aORs against one: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 6: Gender identity and police-recorded offending by offender–victim relationship

	Relationship to victim		
Gender identity	Family member	Other known person	Stranger
Panel A: Adjusted predicted probabilities			
Cisgender men	0.036***	0.040***	0.043***
	(0.001)	(0.001)	(0.001)
Cisgender women	0.015***	0.016***	0.014***
	(0.000)	(0.001)	(0.000)
Transgender men	0.022***	0.018**	0.022**
	(0.008)	(0.007)	(0.009)
Transgender women	0.022***	0.019***	0.017***
	(0.006)	(0.006)	(0.005)
non-binary, assigned male at birth	0.010***	0.027***	0.020***
	(0.004)	(0.006)	(0.005)
non-binary, assigned female at birth	0.007**	0.010***	0.016***
	(0.003)	(0.003)	(0.004)
Differences in adjusted predicted probabilities			
Transgender women − cisgender men	−0.015**	−0.021***	−0.026***
	(0.006)	(0.006)	(0.005)
Transgender women − cisgender women	0.007	0.003	0.003
	(0.006)	(0.006)	(0.005)
Panel B: Adjusted odds ratios			
Cisgender women	0.382***	0.382***	0.298***
	(0.013)	(0.013)	(0.011)
Transgender men	0.581	0.430**	0.493
	(0.228)	(0.175)	(0.215)
Transgender women	0.572*	0.461**	0.361***
	(0.179)	(0.155)	(0.117)
non-binary, assigned male at birth	0.254***	0.665*	0.435***
	(0.097)	(0.145)	(0.115)
non-binary, assigned female at birth	0.180***	0.223***	0.348***
	(0.082)	(0.073)	(0.099)
Transgender women versus cisgender women	1.496	1.206	1.211
	(0.473)	(0.408)	(0.392)
Observations	170,775	170,775	170,775

Notes: Sample: birth cohorts 1993–1998; outcomes at ages 17–26. Outcomes indicate a qualifying record involving the stated relationship. Definitions are given in Section 2.5. Categories may overlap. Panel A reports average adjusted probabilities and differences on the 0–1 scale; differences multiplied by 100 are percentage points. Panel B reports adjusted odds ratios, relative to cisgender men except in the final comparison. Models adjust for year-month of birth, ethnicity, parental-income group, and parental sentencing measures. Postcode-clustered standard errors are in parentheses; aOR standard errors use the delta method. Contrasts use unrounded estimates. Stars test probabilities and differences against zero and aORs against one: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 7: Gender identity and police-recorded offending by offense location

	Police-recorded offense location	
Gender identity	Home or residential	Outside location
Panel A: Adjusted predicted probabilities		
Cisgender men	0.068***	0.083***
	(0.001)	(0.001)
Cisgender women	0.028***	0.042***
	(0.001)	(0.001)
Transgender men	0.051***	0.052***
	(0.012)	(0.013)
Transgender women	0.028***	0.035***
	(0.007)	(0.008)
non-binary, assigned male at birth	0.034***	0.049***
	(0.006)	(0.008)
non-binary, assigned female at birth	0.021***	0.031***
	(0.005)	(0.006)
Differences in adjusted predicted probabilities		
Transgender women − cisgender men	−0.040***	−0.048***
	(0.007)	(0.008)
Transgender women − cisgender women	0.000	−0.007
	(0.007)	(0.008)
Panel B: Adjusted odds ratios		
Cisgender women	0.369***	0.457***
	(0.010)	(0.011)
Transgender men	0.723	0.586*
	(0.184)	(0.160)
Transgender women	0.371***	0.379***
	(0.101)	(0.092)
non-binary, assigned male at birth	0.454***	0.547***
	(0.092)	(0.103)
non-binary, assigned female at birth	0.269***	0.326***
	(0.068)	(0.072)
Transgender women versus cisgender women	1.006	0.828
	(0.271)	(0.200)
Observations	170,775	170,775

Notes: Sample: birth cohorts 1993–1998; outcomes at ages 17–26. Outcomes indicate a qualifying record in a selected home/residential or outside location, as defined in Section 2.5. Outside is not the complement of home, and individuals may appear in both categories. Panel A reports average adjusted probabilities and differences on the 0–1 scale; differences multiplied by 100 are percentage points. Panel B reports adjusted odds ratios, relative to cisgender men except in the final comparison. Models adjust for year-month of birth, ethnicity, parental-income group, and parental sentencing measures. Postcode-clustered standard errors are in parentheses; aOR standard errors use the delta method. Contrasts use unrounded estimates. Stars test probabilities and differences against zero and aORs against one: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Online Appendix

A Additional Tables and Figures

Table A.1: Construction of the analysis sample by gender identity

		Gender identity					
Sample restriction	Total	Cisgender men	Cisgender women	Transgender men	Transgender women	non-binary, assigned male at birth	non-binary, assigned female at birth
Residential population in March 2023	418,785	-	-	-	-	-	-
NZ born via DIA birth records	257,460	-	-	-	-	-	-
Provided self-response on sex at birth and gender identity	206,682	99,201	104,340	378	510	906	1,347
Residential location identified	199,200	95,799	100,401	360	489	855	1,296
Overseas restriction	170,775	83,172	84,954	318	429	762	1,140

Notes: Counts refer to individuals from the 1993–1998 birth cohorts who were resident in New Zealand in March 2023. Restrictions are applied sequentially; each row reports the individuals remaining after the listed and preceding restrictions. New Zealand birth is established through linkage to Department of Internal Affairs (DIA) birth registrations. Gender-identity groups are constructed jointly from self-reported sex at birth and gender in the 2023 Census. Residential location refers to the postcode identified from the address-notification record active 30 days before the individual's 15th birthday. The overseas restriction excludes individuals with 12 or more recorded months overseas at ages 17–26. Overseas exposure is calculated from spells lasting more than 30 days, restricted to their overlap with this age window and summed within individuals. Dashes indicate that gender-specific counts are not reported at that stage.

Table A.2: Gender identity and police-recorded offending: sensitivity to overseas-absence thresholds

	Any selected police-recorded offense, ages 17–26				
	Maximum months overseas				
Gender identity	0	3	6	9	24
Panel A: Adjusted predicted probabilities					
Cisgender men	0.140***	0.130***	0.126***	0.124***	0.119***
	(0.002)	(0.002)	(0.002)	(0.002)	(0.001)
Cisgender women	0.072***	0.066***	0.063***	0.061***	0.059***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Transgender men	0.098***	0.093***	0.087***	0.084***	0.083***
	(0.018)	(0.017)	(0.016)	(0.015)	(0.014)
Transgender women	0.057***	0.051***	0.051***	0.052***	0.052***
	(0.011)	(0.010)	(0.010)	(0.010)	(0.010)
non-binary, assigned male at birth	0.073***	0.067***	0.065***	0.064***	0.062***
	(0.011)	(0.010)	(0.009)	(0.009)	(0.009)
non-binary, assigned female at birth	0.054***	0.050***	0.048***	0.047***	0.045***
	(0.009)	(0.008)	(0.008)	(0.008)	(0.007)
Differences in adjusted predicted probabilities					
Transgender women − cisgender men	−0.083***	−0.079***	−0.075***	−0.072***	−0.067***
	(0.012)	(0.010)	(0.010)	(0.010)	(0.010)
Transgender women − cisgender women	−0.015	−0.015	−0.012	−0.009	−0.007
	(0.011)	(0.010)	(0.010)	(0.010)	(0.010)
Panel B: Adjusted odds ratios					
Cisgender women	0.447***	0.442***	0.434***	0.431***	0.427***
	(0.010)	(0.009)	(0.009)	(0.009)	(0.008)
Transgender men	0.639**	0.658*	0.635**	0.618**	0.645**
	(0.140)	(0.142)	(0.136)	(0.133)	(0.132)
Transgender women	0.339***	0.328***	0.341***	0.357***	0.373***
	(0.077)	(0.074)	(0.073)	(0.075)	(0.078)
non-binary, assigned male at birth	0.453***	0.448***	0.453***	0.454***	0.454***
	(0.075)	(0.073)	(0.074)	(0.073)	(0.074)
non-binary, assigned female at birth	0.317***	0.325***	0.322***	0.320***	0.320***
	(0.061)	(0.059)	(0.058)	(0.057)	(0.055)
Transgender women versus cisgender women	0.760	0.742	0.786	0.828	0.873
	(0.170)	(0.166)	(0.167)	(0.173)	(0.180)
Observations	125,937	146,091	158,589	165,753	182,265

Notes: Sample: New Zealand-born individuals from the 1993–1998 birth cohorts. The outcome indicates any qualifying Police record at ages 17–26. Columns vary the maximum permitted recorded months overseas during this window. Overseas exposure is calculated from spells lasting more than 30 days, restricted to their overlap with ages 17–26 and summed within individuals; the zero-month threshold therefore does not rule out shorter trips. Panel A reports average adjusted probabilities and differences on the 0–1 scale; differences multiplied by 100 are percentage points. Predictions are averaged over the estimation sample in each column. Panel B reports adjusted odds ratios, relative to cisgender men except in the final comparison. Models adjust for year-month of birth, ethnicity, parental-income group, and parental sentencing measures. Postcode-clustered standard errors are in parentheses; aOR standard errors use the delta method. Contrasts use unrounded estimates. Stars test probabilities and differences against zero and aORs against one: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table A.3: Police-recorded offending among gender-minority people AMAB with hormone prescriptions

Gender identity	Any offence, ages 17–26
Panel A: Adjusted predicted probabilities	
Cisgender men	0.122^{***}
	(0.001)
Cisgender women	0.061^{***}
	(0.001)
Gender-minority AMAB with hormone prescriptions	0.054^{***}
	(0.012)
Differences in adjusted predicted probabilities	
Hormone subgroup − cisgender men	-0.069^{***}
	(0.012)
Hormone subgroup − cisgender women	−0.007
	(0.012)
Panel B: Adjusted odds ratios	
Cisgender women	0.431^{***}
	(0.009)
Gender-minority AMAB with hormone prescriptions	0.375^{***}
	(0.094)
Hormone subgroup versus cisgender women	0.871
	(0.218)
Observations	168,468
Gender-minority AMAB with hormone prescriptions	342

Notes: Sample: birth cohorts 1993–1998. The hormone subgroup comprises transgender women and non-binary people AMAB with at least one recorded dispensing of oestrogen and testosterone-blocking medication at ages 17–26. The outcome indicates any qualifying Police-recorded offense in that age window. Panel A reports adjusted probabilities and differences on the 0–1 scale; multiplying differences by 100 gives percentage points. Predictions are averaged over the supplementary estimation sample. Panel B reports adjusted odds ratios, relative to cisgender men except in the final comparison. Adjustment variables are those used in the main analysis. Postcode-clustered standard errors appear in parentheses; aOR standard errors use the delta method. Stars test probabilities and differences against zero and aORs against one: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table A.4: Gender identity and early educational performance

Gender identity	Individuals	Score	p-value
Cisgender men	69,879	0.364	-
Cisgender women	73,188	0.424	<0.001
Transgender men	252	0.393	0.012
Transgender women	330	0.370	0.541
non-binary, assigned male at birth	630	0.404	<0.001
non-binary, assigned female at birth	993	0.488	<0.001

Notes: The sample comprises 145,272 New Zealand-born individuals from the 1993–1998 birth cohorts in the main analysis sample whose first recorded Level 1 expected-percentile score is observed at ages 14–17. Scores coded -1, indicating that no Level 1 standards have been taken, are excluded; scores coded zero, indicating only not-achieved results, are retained. "Individuals" reports the number of individuals in each gender group, and "Score" reports the unadjusted mean score on a 0–1 scale, with higher values indicating stronger educational performance. Reported p-values are from t-tests of equality between each group's mean score and that of cisgender men, who are the reference category in every comparison. The dash denotes the reference group.

Table A.5: Gender identity and police-recorded offending by early educational performance

Gender identity	Any selected police-recorded offense, ages 17–26		
	Available-score sample	Low score	Higher score
Panel A: Adjusted predicted probabilities			
Cisgender men	0.103***	0.182***	0.052***
	(0.002)	(0.003)	(0.001)
Cisgender women	0.050***	0.105***	0.025***
	(0.001)	(0.002)	(0.001)
Transgender men	0.055***	0.101***	—
	(0.014)	(0.031)	
Transgender women	0.046***	0.065***	0.037***
	(0.011)	(0.021)	(0.014)
non-binary, assigned male at birth	0.046***	0.089***	0.024***
	(0.009)	(0.021)	(0.007)
non-binary, assigned female at birth	0.040***	0.114***	0.019***
	(0.007)	(0.026)	(0.005)
Differences in adjusted predicted probabilities			
Transgender women − cisgender men	−0.058***	−0.118***	−0.015
	(0.011)	(0.021)	(0.014)
Transgender women − cisgender women	−0.004	−0.041**	0.013
	(0.011)	(0.021)	(0.014)
Panel B: Adjusted odds ratios			
Cisgender women	0.431***	0.511***	0.445***
	(0.011)	(0.015)	(0.016)
Transgender men	0.484**	0.485**	—
	(0.141)	(0.174)	
Transgender women	0.393***	0.295***	0.695
	(0.105)	(0.103)	(0.275)
non-binary, assigned male at birth	0.401***	0.420***	0.442***
	(0.085)	(0.112)	(0.137)
non-binary, assigned female at birth	0.338***	0.559**	0.345***
	(0.067)	(0.150)	(0.091)
Transgender women versus cisgender women	0.911	0.578	1.562
	(0.242)	(0.201)	(0.613)
Observations	145,272	50,580	94,692

Notes: Sample: New Zealand-born individuals from the 1993–1998 birth cohorts whose first recorded Level 1 expected-percentile score (p_1) is observed at ages 14–17. The outcome indicates any qualifying Police record at ages 17–26. The first column pools individuals with the requisite score information; the other columns report separate regressions by score category: a low score refers to a score belonging to the lowest 3 deciles and a higher score above that threshold. Scores coded −1, indicating that no Level 1 standards have been taken, are excluded; scores coded zero, indicating only not-achieved results, are retained. Panel A reports average adjusted probabilities and differences on the 0–1 scale; differences multiplied by 100 are percentage points. Predictions are averaged over the estimation sample in each column. Panel B reports adjusted odds ratios, relative to cisgender men except in the final comparison. Models adjust for year-month of birth, ethnicity, parental-income group, and parental sentencing measures. Postcode-clustered standard errors are in parentheses; aOR standard errors use the delta method. Dashes denote estimates not reported in the supplied output. Contrasts use unrounded estimates. Stars test probabilities and differences against zero and aORs against one: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.