

# Discrimination Against Immigrants in the Criminal Justice System: Evidence from Pretrial Detentions\*

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## Abstract

This paper tests for discrimination against immigrant defendants in the criminal justice system in Chile using a decade of nationwide administrative records on pretrial detentions. Observational benchmark regressions show that immigrant defendants are 8.6 percentage points less likely to be released pretrial relative to Chilean defendants with similar proxies for pretrial misconduct potential. Diagnostics for omitted variable bias –including a novel test to assess the quality of the proxy vector based on comparisons of pretrial misconduct rates among released defendants– suggest that the discrimination estimates are not driven by omitted variable bias and that, if anything, failing to fully account for differences in misconduct potential leads to an underestimation of discrimination. Our estimates suggest that discrimination stems from an informational problem because judges do not observe criminal records in origin countries, with stereotypes and taste-based discrimination playing a role in the problem’s resolution. We find that discrimination is especially large for drug offenses and that discrimination increased after a recent immigration wave.

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*“When Mexico sends its people, they’re not sending their best. (...). They’re bringing drugs. They’re bringing crime. They’re rapists. And some, I assume, are good people.”*

Donald Trump, presidential announcement speech, June 16, 2015.

## 1 Introduction

Increasing flows of international migrants have induced several policy challenges for destination countries in recent years (UN, 2013, 2019) including, among others, the mitigation of concerns about discrimination (Heath et al., 2013). Labor market, crime, and welfare concerns can trigger negative attitudes towards immigrants, particularly when cultural prejudice and misperceptions regarding foreign born and minority populations are commonplace (Dustmann and Preston, 2007; Hangartner et al., 2019; Grigorieff et al., 2020; Ajzenman et al., 2021; Alesina et al., 2021, 2022; Bursztyrn et al., 2021a). Characterizing discrimination patterns against immigrants is therefore a policy priority to improve the possibilities to integrate immigrants in destination countries.

This paper conducts an empirical analysis of discrimination against immigrant defendants in the criminal justice system, in particular, in the dictation of pretrial detention. Pretrial detention decisions are particularly relevant for two main reasons. First, one of the prevalent narratives that underpins negative attitudes towards immigration is its potential effect on crime (Fasani et al., 2019; Bursztyrn et al., 2021b). Thus, if there is discrimination against immigrants, it should be especially salient in institutions that are concerned with that rubric. Second, the pretrial detention process has been shown to be permeated by discriminatory practices (Arnold et al., 2018, 2021; Grau and Vergara, 2021) and to have a significant impact on the labor market opportunities, conviction probabilities, and the participation in public assistance programs for defendants (Heaton et al., 2017; Leslie and Pope, 2017; Dobbie et al., 2018; Dobbie and Yang, 2021a,b; Grau et al., 2021; Stevenson and Mayson, 2021). These costs are particularly harmful for immigrant populations given the challenges they face relating to integration in general and entry into the labor market in particular (Åslund et al., 2014; Dustmann et al., 2017; Brell et al., 2020; Egger et al., 2022).

We study this scenario in Chile where the pretrial detention process functions in a similar manner to the US—that is, judges decide on the pretrial detention status of defendants using limited information and in a limited time period with a mandate to not detain defendants unless misconduct potential (pretrial recidivism or failure to appear in court, or both) is considerable. Unlike in the US, however, the Chilean system is uniform across all localities and does not permit monetary bail. Since approximately 2013 Chile has become an important destination for immigrants

from other Latin American countries. And, as we discuss in Section 2, this immigration flow has generated hostile discourses in the public sphere and misperceptions regarding immigrants and immigration in the general population.

We conduct the empirical analysis using nationwide administrative records for the period 2008—2017. Our database covers 95% of Chile’s criminal prosecutions and includes detailed information on defendants and prosecutions. Despite having cleaner criminal records upon prosecution and lower pretrial misconduct rates when released, immigrant defendants are more likely to be detained pretrial relative to defendants who are Chilean nationals. We formally explore whether this descriptive fact constitutes some form of aggregate discrimination by estimating observational benchmark regressions—that is, by testing if defendants with similar proxies for misconduct potential face, on average, different pretrial release rates because of their immigration status, an exercise similar in spirit to [Gelman et al. \(2007\)](#), [Abrams et al. \(2012\)](#), and [Rehavi and Starr \(2014\)](#). The main result shows that, conditional on type of crime, criminal records, judge leniency, attorney quality, and court-by-time fixed effects, immigrant defendants are 8.6 percentage points (pp) more likely to be detained pretrial relative to Chilean defendants, a difference that represents 53% of the unconditional baseline pretrial detention rate.

We perform two sets of diagnostics that suggest that our main estimates are not driven by systematic differences in unobserved misconduct potential (i.e., omitted variable bias, OVB). First, we implement the method set out in [Oster \(2019\)](#) for assessing the degree of selection on unobservables using observed covariates. We conclude that, if anything, our observational benchmark regression underestimates discrimination against immigrants. Specifically, we show that the discrimination estimate is attenuated (although not reversed) only if we impose the nonstandard assumption that unobservables are negatively correlated with the control variables—that is, if defendants with observables that induce lower release rates are also endowed with unobservables that work in the opposite direction. This is consistent with the fact that immigrants have, on average, cleaner criminal records. In fact, and contrary to the common finding in observational benchmark regressions, we find that the release disparity between Chilean defendants and immigrant defendants is 30% smaller (6.6 pp) when control variables are omitted.

The second diagnostic is a novel test that uses observed pretrial misconduct rates of released defendants to assess the degree of OVB in the release equation. The intuition is as follows: OVB arises if there are unobservables that correlate with immigration status and predict misconduct potential. Then, we can use observed pretrial misconduct of released defendants to test whether the included control variables are sufficient to account for systematic differences in pretrial misconduct

rates between immigrant defendants and Chilean defendants. Because the sample of released defendants is selected, we estimate different selection models using assigned judges and attorneys as the excluded variables and show that, conditional on type of crime, criminal records, and time-varying court characteristics, immigration status has no predictive power on observed pretrial misconduct. This confirms that the set of controls included in the benchmark regression is sufficient to account for systematic risk differences between Chilean defendants and immigrant defendants.

The second test also suggests that, while time-varying court characteristics matter for the discrimination point estimate, they do little to correct for OVB and, therefore, they may induce included variable bias (IVB) in the main estimations. In other words, rather than controlling for differences in unobserved misconduct potential, court-by-time variation seems to mediate discrimination through the systematic assignment of immigrant defendants to less lenient courts. Under this interpretation, the baseline discrimination estimate of 8.6 pp constitutes a conservative bound given that excluding the court-by-time fixed effects increases the release disparity gap to 10.3 pp.

After establishing the robustness of the reduced-form discrimination estimate, we perform different heterogeneity analyses to obtain additional insights about the aggregate pattern. We first explore whether discrimination against immigrants stems from an informational problem. The criminal records of immigrant defendants are likely to be censored because judges cannot observe the records kept in the country of origin. Then, judges may process the information contained in criminal records differently depending on whether the defendant is immigrant or not. To test this hypothesis, we estimate our main specification separately for samples of defendants with and samples of defendants without previous prosecutions in Chile. We find that the conditional release disparities are 1.9 pp and 13.2 pp for defendants with and defendants without previous prosecutions, respectively. This implies that, unlike Chilean defendants, immigrant defendants are not rewarded for having clean criminal records. In fact, comparing the discrimination estimates with the unconditional differences in release rates suggest that judges systematically impute nonzero criminal histories for all immigrant defendants, regardless of their criminal records in Chile. We complement this analysis by showing that the discrimination estimate is 36% smaller for experienced judges. Both pieces of evidence strongly suggest that the informational problem seems to be the main driver behind the discrimination patterns.

The fact that discrimination stems from an informational problem does not necessarily mean that the aggregate estimate can be rationalized by accurate statistical discrimination. On one hand, the informational problem can be circumvented based on stereotypes (biased beliefs). On the other hand, the informational problem may provide the room for judges to exercise preferences

against immigrants (taste-based discrimination). To obtain further insights, we estimate observational outcome tests (Knowles et al., 2001; Grau and Vergara, 2021) that, when rejected, provide evidence of a nonseparately identified combination of stereotypes and taste-based discrimination (Hull, 2021). The outcome test is rejected suggesting that the discrimination pattern, although stemming from an informational problem, cannot be rationalized by accurate statistical discrimination. The combination of stereotypes and taste-based discrimination seems to play an important role in explaining the aggregate estimates of discrimination against immigrants.

We also explore whether or not the patterns of discrimination vary according to crime category. We find that discrimination is particularly severe for drug offenses: conditional release disparities are around four times larger than the baseline benchmark estimate. Statistically and economically significant differences are also found for thefts and robberies and other property crimes. Other violent crimes like homicides and sexual offenses show no significant differences, although the number of immigrants that are charged with these types of crimes is too small to confidently detect an effect. Crime categories that are associated with less hostile discourses in the public sphere (such as white collar crimes or crimes against privacy) also display no significant release disparities. This heterogeneity is consistent with the discourse that links immigration flows to drug trafficking and other crimes that tend to be prominent in surveys that measure citizens perceived fears of crime playing a role in affecting preferences and stereotypes.

Finally, we explore the evolution of discrimination patterns during the recent wave of immigration to Chile. Immigration inflows in Chile grew considerably and became more racially diverse after 2012. We find that benchmark estimates are 55% larger in the 2013–2017 period relative to the 2008–2012 period. We also find that discrimination estimates are remarkably similar across different groups of immigrants, suggesting that the increase over time is driven by an intensification of discrimination rather than a composition effect. Although we do not establish a causal relationship, we conjecture that this trend is related to the crime-related narrative response to the immigration wave lead by political authorities and the media (Ajzenman et al., 2021).

## 2 Setting and Data

This section describes the Chilean context and institutional setting. We start by characterizing immigration in Chile, discussing its magnitude, trends, composition, and the attitudes to immigration among Chilean citizens. We then describe the pretrial detention system. Finally, we describe the data used in the empirical analysis and provide descriptive statistics.

## 2.1 Immigration in Chile

Immigration in Chile has changed dramatically since 2013. According to the national census, immigrants represented around 1.5% of the Chilean population in 2002 and in 2012 this figure was 2%. In 2017, however, the share of immigrants in the population increased to 4.5%, and projections from the National Institute of Statistics suggest that the number reached 7.8% in 2019. In terms of gross inflows, both the volume and composition has changed considerably. Between 2008 and 2012 the inflow of immigrants was fairly stable at around 100,000 individuals per year. Peruvians accounted for the largest share of immigrants arriving each year (roughly 50% of new arrivals) and Argentinians, Bolivians, and Colombians collectively represented around 25% of inflows. The annual inflow of immigrants grew to around 350,000 individuals in 2017. In addition, the composition changed substantially. For the 2008–2012 period, Haitian and Venezuelan immigrants accounted for around 2% of the arrivals; however, for the 2013–2017 period they represented almost 20% of the arrivals.<sup>1</sup> Inflows of Colombian nationals also increased significantly after 2013. Their share among new arrivals to Chile grew from 10% in the period 2008–2012 to 19% in the period 2013–2017.

The recent immigration wave has not gone unnoticed by Chile’s population, politicians, and media. Similar to the inhabitants of other countries (Hopkins et al., 2019; Jørgensen and Osmundsen, 2020; Alesina et al., 2022), Chileans overestimate the size of the immigrant population. As summarized in Ajzenman et al. (2021), surveys show that in 2019 Chileans thought that immigrants represented 33% of the total population, which compares unfavorably to the administrative estimate of 7.8%. Additional survey evidence shows that *increase in crime* is the most prominent concern related to immigration among Chile’s population. Ajzenman et al. (2021) report that these perceptions are not related to changes in victimization rates, even though the concerns were severe enough to induce increases in preventive behavioral responses, such as investments in home security. In addition, Ajzenman et al. (2021) show that the media plays an important role in the divergence between perceptions and actual crime rates, for example, by increasing the number of crime-related news segments with an alleged perpetrator that is an immigrant.

Public attitudes toward immigrants can also be inferred from recent political campaigns. Immigration, which was not a public issue before the recent increase in immigration inflows, has become a prominent topic in the national political debate, with an increasing number of politicians framing their public discourse in the context of immigration regulation. For example, inspired by

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<sup>1</sup>To illustrate the prominence of Chile as a destination of choice within the region, in March 2021 Chile was the country with the second highest number of Venezuelan immigrants equating to 600,000 individuals with regular status, just below Colombia with 720,000. See <https://data2.unhcr.org/en/situations/vensit> for details.

Donald Trump, the far-right candidate in the last presidential election, José Kast, proposed to build a ditch on Chile’s border with Bolivia. In most cases, a link between crime and immigration is proffered as the most important reason for stopping or reducing the arrival of new immigrants. For example, when the National Congress of Chile passed President Piñera’s migration project in 2019, he described the initiative as a “(...) great step forward to continue putting our house in order regarding matters of migration, and thus better combat illegal immigration and the evils such as delinquency, drug trafficking, and organized crime entering Chile” (Twitter, January 16, 2019). These types of statements, that display either implicit or explicit anti-immigrant sentiment, have been accompanied by instances of misinformation that appear to be designed to stigmatize immigrants or immigration. For example, in 2018, a widely shared social media post claimed that the former President Bachelet received USD 3,000 from the UN for each Haitian immigrant that entered Chile. The increase in negative attitudes towards immigrants has contributed towards Chile’s political climate becoming more polarized and an increase in the relative popularity of far-right nationalist groups with conservative-leaning voters.<sup>2</sup>

## 2.2 Pretrial Detentions in Chile

The procedure to define pretrial detention for people arrested in Chile works as follows: During the 24 hours after the initial detention, there is an arraignment hearing in which a detention judge must decide whether the defendant will be incarcerated during the criminal investigation. Monetary bail is not an option in the Chilean system. Following the legal principle of the presumption of innocence, judges should not incarcerate defendants unless there is clear danger of escape (i.e., a high probability of failing to appear in court), the defendant represents a significant threat to society (i.e., a high probability of committing another crime during the investigation), or the imprisonment of the defendant aids the investigation of the criminal case. As described in [Grau et al. \(2021\)](#), pretrial detention became more frequent between 2007 (17,891 cases) and 2018 (34,815 cases). In the same period, the proportion of cases where pretrial detention is imposed has increased from

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<sup>2</sup>There is a contrast between the hostile expressions against immigrants, such as President’s Piñera’s tweet, with what is usually presented as the traditional view on Chilean attitudes towards immigration as expressed in Chito Faró’s famous 1942 folk song “Si vas para Chile” (“If you go to Chile”). This popular and traditional song describes Chileans as particularly receptive to foreigners stating: “Y verás como quieren en Chile al amigo cuando es forastero” (“and you will see how they love a friend in Chile when is a foreigner”). Other music artists have offered a different perspective on attitudes towards immigration. For example, a 2012 track by Chilean rapper Portavoz called “Poblador del mundo” (“World citizen”) tells a story in the voice of an immigrant that feels constant discrimination living in Chile. In one verse, he raps “Llegué con fe para doblarle la mano al destino y me encontré con que aquí soy un enemigo. Y es que con mi llegada los dramas también han venido, ya me han hecho entender que no soy para nada bienvenido” (“I arrived with the belief that I could turn the hand of destiny but found that here I am an enemy. With my arrival the dramas have also come, they have already made me understand that I am not welcome at all”).

7.3% to 9.6%, and pretrial detainees as a share of total prisoners rose from 21.9% to 36%.

Pretrial detention hearings are brief (lasting about 15 minutes) and judges have to make a decision whether or not to detain the defendant at the end of the hearing. Judges are assigned in a quasi-random fashion at the court-by-time level: in every court at the beginning of each month judges are assigned to different time slots to lead arraignment hearings for no reason other than splitting this duty among the judges in the particular court. The assignment of public attorneys to each case is also quasi-random at the court-by-time level.

## 2.3 Data

We use administrative records from the Public Defender’s Office (PDO). The PDO is a centralized public service under the oversight of the Ministry of Justice. It offers criminal defense services to all individuals accused of or charged with a crime; as such, it ensures the right to a defense by a lawyer and due process in criminal trials. Our estimation sample covers more than 95% of the criminal cases for the period between 2008 and 2017, and contains detailed case and defendant characteristics, including the nationalities of the defendants. We define immigrant defendants as defendants that are nationals of any other country than Chile. In addition, we can identify the judges assigned to each case at the beginning of the criminal process (i.e., when pretrial detention decisions take place), as well as the public attorney assigned to the arraignment hearing.

To build the estimation sample, we consider all detention hearings for adult defendants who were arrested between 2008 and 2017 and defended by the PDO. We exclude hearings due to legal summons because the information set available to the judge may be different in those cases. To be able to focus on arraignment hearings in which pretrial detention is a plausible outcome, we only consider types of crimes that have at least a 5% probability of pretrial detention. For the same reason, when defendants are accused of more than one crime during the same arraignment hearing, we only retain the information related to the most severe crime (with severity measured as the probability of pretrial detention). A more detailed description of the data, the sample restrictions, and the variables is presented in Appendix B.

Table 1 shows descriptive statistics for the estimation sample. This table provides three stylized facts that motivate our research question. First, relative to Chilean defendants, immigrants are more likely to be detained before the trial. The pretrial release rate for immigrant defendants and Chilean defendants is 77% and 84%, respectively. Second, immigrant defendants have cleaner criminal records. Chilean defendants are more likely to have previous prosecutions, convictions,

Table 1: Descriptive Statistics

	Chilean	Immigrant
Released	0.84	0.77
<b>Outcomes (only for released)</b>		
Nonappearance in court	0.17	0.16
Pretrial recidivism	0.19	0.13
Pretrial misconduct	0.29	0.24
<b>Individual Characteristics</b>		
Male	0.88	0.87
At least one previous case	0.68	0.41
At least one previous pretrial misconduct	0.40	0.22
At least one previous conviction	0.65	0.39
No. of previous cases	4.58	2.48
Severity previous case	0.09	0.05
Severity current case	0.18	0.18
<b>Court Characteristics</b>		
Average severity (year/Court)	0.09	0.11
No. of cases (year/Court)	3,023	3,557
No. of judges (year/Court)	46	58
<b>Observations (released)</b>	580,400	4,900
<b>Observations (nonreleased)</b>	112,606	1,462

**Notes:** This table presents descriptive statistics for our estimation sample. The sample considers all arraignment hearings for adult defendants who were arrested between 2008 and 2017. We do not include hearings due to legal summons and only consider types of crimes with at least a 5% probability of pretrial detention. When defendants are accused of more than one crime, we retain the information related to the most severe crime (with severity measured as the probability of pretrial detention). More details about the data and the variables can be found in Appendix B.

and instances of pretrial misconduct. The criminal records of immigrants may be censored because judges cannot observe the criminal histories of immigrant defendants before they arrive to Chile. Therefore, how judges interpret immigrants' (lack of) criminal records could play a role in explaining the release disparities. Third, once released, immigrants are less likely to be engaged in any kind of pretrial misconduct; pretrial recidivism rates are 6 pp larger for Chilean defendants.

### 3 Theoretical Framework

To motivate a formal test for discrimination in pretrial detention decisions, this section develops a simple theoretical framework to guide the empirical analysis and formally defines the notion of discrimination we use throughout the paper.

### 3.1 Model

Let  $i$  index defendants and  $j$  index judges, with judges assigned to defendants according to the mapping  $j(i)$ . Let  $Y_i^* \in \{0, 1\}$  be pretrial misconduct if released, with  $Y_i$  being the realized value, so  $Y_i = Y_i^*$  for pretrial released defendants and  $Y_i = 0$  for pretrial detained defendants. Judges are mandated to detain defendants with  $Y_i^* = 1$  and release defendants with  $Y_i^* = 0$  but they do not observe  $Y_i^*$  at the time of the decision.<sup>3</sup> Instead, judges base the release decision on their predictions on the probability that  $Y_i^* = 1$ . The prediction is made using all the available information, which includes the nationality of the defendant,  $I_i = 1\{i \text{ is immigrant}\}$ , and a vector of other characteristics,  $Z_i$ .

The principle of the presumption of innocence and the documented impact of pretrial detention imply that only defendants with a high predicted probability of pretrial misconduct should be detained. Because the decision is discretionary and judges potentially have heterogeneous preferences, release thresholds are idiosyncratic. Then, the release equation for defendant  $i$  can be written as

$$Release_i = 1 \{p_{j(i)}(I_i, Z_i) \leq t_{j(i)}(I_i, Z_i)\}, \quad (1)$$

where  $p_{j(i)}(I_i, Z_i) \in [0, 1]$  is the prediction of the pretrial misconduct probability of defendant  $i$  made by the assigned judge  $j(i)$ , and  $t_{j(i)}(I_i, Z_i) \in [0, 1]$  is the judge-specific release threshold.

Let  $p(I_i, Z_i) = \Pr(Y_i^* = 1 | I_i, Z_i)$ . Because the conditional distribution is possibly unknown to judges, their predictions regarding pretrial misconduct are potentially biased. Let  $b_{j(i)}(I_i, Z_i)$  be the judge-specific bias in predictions, such that  $p_{j(i)}(I_i, Z_i) = p(I_i, Z_i) + b_{j(i)}(I_i, Z_i)$ . Then,

$$Release_i = 1 \{p(I_i, Z_i) \leq t_{j(i)}(I_i, Z_i) - b_{j(i)}(I_i, Z_i)\} \equiv 1 \{f_{j(i)}(I_i, Z_i) \geq 0\} = R(I_i, Z_i, j(i)). \quad (2)$$

As [Hull \(2021\)](#) stresses, this analysis does not require additional assumptions on judge behavior, relating to the way they set thresholds or make predictions, for example. The analysis only needs the judge decision to be binary and to follow a threshold-crossing model. As discussed in [Section 2](#), both assumptions represent a good approximation of the Chilean setting. In this simple model, the decision is deterministic from the judge's perspective. We denote the function that determines whether a defendant with immigration status  $I_i$  and characteristics  $Z_i$  is released or not when assigned to judge  $j(i)$  by  $R(I_i, Z_i, j(i)) \in \{0, 1\}$ .

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<sup>3</sup>Alternatively,  $Y_i^*$  could be defined as the probability of defendants engaging in some form of pretrial misconduct if released, with  $Y_i^* \in [0, 1]$  and the judges' mandate to detain defendants pretrial when  $Y_i^*$  exceeds some threshold.

### 3.2 Sources of Release Disparities and Normative Definitions

Equation (2) suggests that a defendant’s nationality may affect the release decision through different channels. First, judges may use  $I_i$  to predict (or interpret signals of) variables they do not observe that affect  $Y_i^*$  if distributions vary with  $I_i$ . This is defined as statistical discrimination (Aigner and Cain, 1977) and directly affects the conditional outcome probabilities,  $p(I_i, Z_i)$ , and the bias,  $b_{j(i)}(I_i, Z_i)$ , if these parameters impact how priors are updated. Second, if statistical discrimination is inaccurate and the degree of inaccuracy varies with  $I_i$ , this induces a differential prediction bias to the decision. We refer to this as biased beliefs or stereotypes, and its effect is captured by  $b_{j(i)}(I_i, Z_i)$  (Bordalo et al., 2016; Bohren et al., 2021). Third, release thresholds may depend on  $I_i$  because of animus. This affects  $t_{j(i)}(I_i, Z_i)$  and gives form to the standard notion of taste-based discrimination or discriminatory preferences (Becker, 1957, 1993).

In this paper, we focus on an aggregate notion of discrimination that incorporates the three sources of release disparities mentioned: statistical discrimination, biased beliefs or stereotypes, and taste-based discrimination. A possibly complex combination of these drivers of release disparities may yield a reduced-form effect of  $I_i$  on  $Release_i$ , captured by the function  $f_{j(i)}(I_i, Z_i)$  even within defendants with equal  $Y_i^*$ . As emphasized in Yang and Dobbie (2020) and Arnold et al. (2021), all sources of disparities are problematic through the lens of legal principles similar to the Equal Protection Clause in the US Constitution, which suggests that this reduced form effect has normative content regardless of the specific sources. Intuitively, the composite effect of all sources may lead defendants with equal misconduct potential but different immigration status to face different pretrial detention rates. This motivates the following definition.

DEFINITION (DISCRIMINATION): *discrimination,  $\mathbf{D}$ , is defined as*

$$\mathbf{D} = \mathbb{E}[\mathbf{d}(Y_i^*)] = \mathbb{E}[\mathbb{E}[d(Z_i, j(i))|Y_i^*]] = \mathbb{E}[\mathbb{E}[R(1, Z_i, j(i)) - R(0, Z_i, j(i))|Y_i^*]], \quad (3)$$

where the inner expectation integrates over  $(Z_i, j(i))$ , and the outer expectation integrates over  $Y_i^*$ .

Under this discrimination definition, there is system-wide discrimination against immigrants if  $\mathbf{D} < 0$ . The inner object,  $d(Z_i, j(i))$ , is the judge-specific gap in release rates between immigrant and native defendants endowed with  $Z_i$ .  $\mathbf{d}(Y_i^*)$  is the expectation of the judge-specific measure over  $(Z_i, j(i))$  conditional on misconduct potential  $Y_i^*$  and, therefore, does not impose restrictions on the group-specific distributions. Finally, the aggregate measure of discrimination averages over the distribution of misconduct potential. Similar to the definition outlined in Arnold et al. (2021), this definition compares release rates between immigrant and native defendants with equal pretrial

misconduct potential, with potential differences arising because of statistical discrimination, biased beliefs, or taste-based discrimination, or a combination of these sources of discrimination.

**Empirical test** One fundamental challenge when testing for discrimination following (3) is that  $Y_i^*$  is most likely not observable; therefore, a regression of  $Release_i$  on  $I_i$  controlling for  $Y_i^*$  is potentially not implementable. Arnold et al. (2021) propose a weighting scheme that equalizes group-specific distributions of  $Y_i^*$  under which omissions of  $Y_i^*$  from the (weighted) regression of  $Release_i$  on  $I_i$  do not induce OVB and, therefore, identifies  $\mathbf{D}$ . Unfortunately, the immigrant population in Chile is small, which prevents us from implementing this methodology.<sup>4</sup> As an alternative, we implement observational benchmark regressions similar in spirit to Gelman et al. (2007), Abrams et al. (2012), and Rehavi and Starr (2014). Formally, we assume that  $Y_i^*$  can be expressed as a function of observed variables,  $X_i$ , so the discrimination definition can be written as

$$\mathbf{D} = \mathbb{E}[\mathbf{d}(g(X_i))] = \mathbb{E}[\mathbb{E}[d(Z_i, j(i))|g(X_i)]] = \mathbb{E}[\mathbb{E}[R(1, Z_i, j(i)) - R(0, Z_i, j(i))|g(X_i)]], \quad (4)$$

where  $Y_i^* = g(X_i)$ . Although  $X_i$  and  $Z_i$  possibly overlap,  $X_i$  does not have to be equal to  $Z_i$ . Under this assumption, regressing  $Release_i$  on  $I_i$  controlling for  $g(X_i)$  identifies  $\mathbf{D}$ .

In Section 4 we discuss our choice of  $X_i$  and provide diagnostics to assess its performance in approximating  $Y_i^*$ . This is central to the analysis because the extent to which (4) constitutes a good approximation of (3) fundamentally depends on the vector  $X_i$ . More specifically, as noted by Arnold et al. (2021), observational benchmark regressions can suffer both from IVB and OVB. IVB arises when some of the included variables in  $X_i$  mediate discrimination against, in this case, immigrant defendants in a disparate impact sense. For example, if there is discrimination based on place of living and immigrants are overrepresented in a discriminated locality, then controlling for place of living will attenuate the estimated disparity. OVB arises when the vector  $X_i$  does not fully account for group differences in misconduct potential and, therefore, the estimated coefficient reflects differences in unobservables rather than discriminatory forces. This empirical challenge generates a tension between over and under controlling.

We deal with this tension by choosing the smallest set of variables that successfully eliminates

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<sup>4</sup>The Arnold et al. (2021) method combines quasi-random assignment of judges, observational misconduct data on released defendants behavior, and extrapolation techniques. The extrapolation step requires several judges handling several cases that involve defendants that are immigrants. Our limitation is, therefore, statistical power. Specifically, the extrapolation technique used in Arnold et al. (2021) is based on regressing average pretrial misconduct rates of released immigrant defendants against judge leniency to infer the average pretrial misconduct rate of a supremely lenient judge using an identification at infinity argument. For this technique to work properly several judges are required in order to have enough support, in addition to several cases per judge so that pretrial misconduct rates can be precisely estimated. Both requirements are not met in our immigrant subsample.

OVB and therefore minimizes the likelihood of IVB. Section 4 presents several tests that reject OVB given our choice of  $X_i$  and develops sensitivity analyses that flag potential sources of IVB in our setting. In particular, variables that do little to attenuate OVB concerns and significantly affect the discrimination estimate are more likely to contaminate the result with IVB.

**On the normative interpretation** Related to the discussion on IVB and OVB, our definition of discrimination seeks to identify a disparate impact (rather than disparate treatment) notion of discrimination (Arnold et al., 2021; Rose, 2022). The disparate impact doctrine focuses on release disparities among defendants with equal misconduct potential but different group membership. Under this perspective, different structural models that rationalize similar release gaps conditional on misconduct potential are normatively equivalent. On the other hand, the disparate treatment doctrine focus on the actual release decision, understanding as discrimination explicit judge-level practices against a particular group. This implies that the normative implication of release disparities among defendants with equal misconduct potential is different if they follow explicit discriminatory practices against group membership or if, for example, are driven by discriminatory practices against other characteristics that correlate with group membership.

As before, the extent to which our definition relates to each normative perspective depends on the variables included in  $X_i$ . More specifically, it is dependent on how accurately these variables proxy misconduct potential. If  $X_i$  exactly matches the information set of judges ( $Z_i$ ), the definition will probably approximate a disparate treatment notion of discrimination because judges observe variables through which discrimination towards immigrants manifests itself and that are unlikely to predict misconduct potential on top of other controls. That is why we purposely exclude (and why we are not concerned that we do not observe) all demographics other than immigration status that are potentially observed by judges (age, place of living, physical aspects, or income, for example). By contrast, we include the minimum set of variables that, we argue, are sufficient to eliminate OVB because they constitute a reasonable proxy of misconduct potential—namely, criminal records and case characteristics.<sup>5</sup>

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<sup>5</sup>Arnold et al. (2021) show that discrimination estimates are biased when criminal record and type of crime variables are included in regressions that already control for misconduct potential through an IVB mechanism, which implies that these variables are imperfect proxies for misconduct potential. This can be explained if, for example, the distributions of types of crime differ by immigration status and judges are engaged in crime-specific discrimination patterns, or if there is discrimination against immigrants in previous stages of the judicial process that is manifest in differences in the criminal records. Note, however, that Arnold et al. (2021) estimated observational release disparities that control for type of crime and criminal records are close in magnitude to the weighted benchmark regressions that recover the unbiased estimate, which suggests that this set of controls do a reasonably good job when approximating potential misconduct. They document unconditional gaps of 7.2 pp that are reduced to 5.2 pp when including controls and fixed effects. The bias-corrected weighted regressions with no controls report discrimination estimates that range from 4.2 pp to 5.4 pp, depending on the specific extrapolation technique.

## 4 Empirical Analysis

This section presents the empirical strategy and the main results. We estimate substantial discrimination against immigrant defendants in pretrial detention decisions. We discuss and implement different tests for assessing the degree of OVB and conclude that unobservables are unlikely to explain the results. If anything, unobserved differences in misconduct potential are likely to exert downward bias on the discrimination estimates.

### 4.1 Benchmark Regressions

**Empirical strategy** Discrimination, as defined by equation (4), can be potentially identified by *benchmark regressions*—that is, by reduced-form regressions of  $Release_i$  on  $I_i$  and  $g(X_i)$ . We assume that  $g$  is linear in  $X_i$ —that is,  $g(X_i) = X_i' \alpha_X$ . And, importantly for the discussion on OVB, we further define  $X_i = [X_i^o \ X_i^u]$ , where  $o$  accounts for variables that are observed by the econometrician and  $u$  accounts for variables that are unobserved by the econometrician. The observational benchmark test therefore consists of regressions of the following type:

$$Release_i = \alpha_0 + \alpha_D I_i + X_i^{o'} \alpha_{X_o} + \varepsilon_i, \quad (5)$$

where  $\varepsilon_i = X_i^{u'} \alpha_{X_u} + \xi_i$  is the error term. We distinguish between two sources of unobserved heterogeneity that affect the release decision. First,  $X_i^{u'} \alpha_{X_u}$  accounts for unobservables that predict misconduct potential. The vector  $X_i$  that accurately predicts  $Y_i^*$  may be extensive and may therefore contain information that is not observed by the econometrician. Second,  $\xi_i$  accounts for unobservables that do not predict misconduct potential but may affect the release decision. For example,  $\xi_i$  may contain variables that judges use to discriminate on top of immigration status but that are not correlated with  $Y_i^*$  after controlling for  $X_i$ .

Let  $\hat{\alpha}_D$  be the ordinary least squares (OLS) estimator of (5). The main identification assumption for  $\mathbb{E}[\hat{\alpha}_D | I_i, X_i^o] = \mathbf{D}$  is  $\mathbb{C}(I_i, X_i^u) = 0$ . That is to say, the unobservables that predict misconduct potential,  $X_i^u$ , are not correlated with immigration status,  $I_i$ . This follows from equation (4) because discrimination is defined as average release disparities conditional on misconduct potential. If  $I_i$  and  $X_i^u$  are correlated, then  $\hat{\alpha}_D$  will capture both discrimination and unobserved differences in misconduct potential and will therefore be contaminated by standard OVB. Note, however, that correlation between  $I_i$  and  $\xi_i$  is allowed by our definition of discrimination given our focus on a disparate impact notion of discrimination so its presence is, therefore, not problematic for the

Table 2: Main Results–Benchmark Regressions:  $\hat{\alpha}_D$ 

	(1)	(2)	(3)	(4)	(5)
Immigrant	-0.067 (0.014)	-0.103 (0.013)	-0.063 (0.013)	-0.030 (0.012)	-0.087 (0.012)
Mean dep. variable	0.84	0.84	0.84	0.84	0.84
Individual controls	No	Yes	No	No	Yes
Attorney and judge controls	No	No	Yes	No	Yes
Court-by-year fixed effects	No	No	No	Yes	Yes
No. of Immigrants	6,362	6,362	6,362	6,362	6,362
No. of Chileans	693,006	693,006	693,006	693,006	693,006
R-squared	0.000	0.167	0.011	0.023	0.191

**Notes:** This table presents the estimated benchmark coefficient,  $\hat{\alpha}_D$ , of equation (5) and its corresponding standard error in parentheses (clustered at the court-by-year level). Each column represents a different regression that includes different sets of controls. *Individual controls* include an indicator for whether the individual has previous prosecutions, the number of previous prosecutions, the severity of the last prosecution (measured as the average pretrial detention rate of the type of crime), an indicator for whether the individual was engaged in pretrial misconduct during a previous prosecution, an indicator for whether the individual has been convicted of a crime in the past, and type of crime fixed effects for the current prosecution (consisting of nine mutually exclusive categories). *Attorney and judge controls* include judge leniency and public attorney quality and their squares, measured as residualized (against court-by-year fixed effects) leave out release rates, as in [Dobbie et al. \(2018\)](#).

identification of  $\mathbf{D}$ . After presenting the main results, we discuss and implement diagnostics to assess the pervasiveness of potential OVB in our main benchmark estimations.

We estimate equation (5) using OLS and include the following sets of controls in the vector  $X_i^o$ . First, we include individual-level characteristics related to the criminal history and the type of crime. These include an indicator for whether the individual has previous prosecutions, the number of previous prosecutions, the severity of the last prosecution (measured as the average pretrial detention rate of the type of crime), an indicator for whether the individual was engaged in pretrial misconduct during a previous prosecution, an indicator for whether the individual has been convicted of a crime in the past, and type of crime fixed effects for the current prosecution. These variables are expected to proxy individual-level misconduct potential. Second, to improve precision, we include measures of judge leniency and public attorney quality and their squares, measured as residualized (against court-by-year fixed effects) leave out release rates, as in [Dobbie et al. \(2018\)](#). Third, we include court-by-year fixed effects that can potentially account for unobserved shocks that correlate with misconduct potential. Standard errors are clustered at the court-by-year level.

**Results** Table 2 presents the results. Each column represents a different regression and reports the estimated benchmark coefficient,  $\hat{\alpha}_D$ , with its corresponding standard error. In the absence

of OVB,  $\hat{\alpha}_D$  is a valid measure of discrimination, with negative values accounting for discrimination against immigrant defendants. Column 1 shows that the raw (uncontrolled) release rate for immigrants is 6.7 pp lower than the release rate for defendants that are Chilean nationals. The difference is quantitatively important: it represents 8% of the unconditional average release rate or, put differently, almost 42% of the unconditional average detention rate.

If there are differences in misconduct potential between immigrant defendants and Chilean defendants, this result could be explained by OVB. Columns 2, 3, and 4 explore how the gap changes when adding the specific sets of controls previously discussed. When adding the individual controls (Column 2), the release rate gap increases to 10.3 pp. Under the assumption that criminal records and type of crime are positively correlated with misconduct potential, this suggests that the raw gap cannot be explained by differences in latent risk. In fact, the result conjectures that ignoring differences in misconduct potential may lead to an underestimation of the measure of discrimination. When adding the (residualized) controls for the assigned judge and attorney (Column 3), the release disparity remains unchanged. This is consistent with judges and attorneys being quasi-randomly assigned at the court-by-year level. When adding court-by-year fixed effects (Column 4), the release rate gap decreases to 3 pp. There are two potential explanations for this: The first explanation is that court-by-time fixed effects could mediate discrimination against immigrant defendants if they are systematically assigned to more severe courts (i.e., IVB). The second explanation is that the 3 pp decrease captures unobserved heterogeneity in misconduct potential that varies at the court-by-year level. Below we present suggestive evidence in favor of the first interpretation. Finally, when adding the three sets of controls (Column (5)), the gap in release rates between immigrant defendants and Chilean defendants is 8.7 pp. This represents 10.3% of the unconditional average release rate or, put differently, 54.3% of the unconditional average detention rate. We interpret these results as strong evidence of discrimination against defendants that are immigrants.

## 4.2 Assessing Omitted Variable Bias

The main threat to identification in benchmark regressions is OVB. If there are omitted variables that predict misconduct potential and that are correlated with immigration status, then the discrimination estimate will be biased. We conduct two sets of diagnostics to assess the extent to which OVB affects our main estimations. Both exercises suggest that unobservables that predict misconduct potential are unlikely to explain the results and, if anything, they are likely to exert downward bias on our discrimination estimates. The second diagnostic also provides elements for assessing the degree of IVB in our benchmark regressions.

**Diagnostic 1: Bounds on selection on unobservables** One way of assessing the role of unobservables is to follow the extension proposed by [Oster \(2019\)](#) to the method proposed in [Altonji et al. \(2005\)](#). This approach interprets observables as unobservables that the econometrician happened to observe, establishing an implicit relationship between them. Under this interpretation, and assuming some degree of correlation between  $X_i^o$  and  $X_i^u$ , the sensitivity of  $\hat{\alpha}_D$  to the inclusion of observed control variables is informative about potential OVB because the bias arising from omitting observables is related to the bias induced by the unobserved variables. This allows for the computation of bounds on  $\hat{\alpha}_D$  given different assumptions on the informativeness of observables about unobservables.

Based on (5), let  $W_{1i} = X_i^o \alpha_{X_o}$  and  $W_{2i} = X_i^u \alpha_{X_u}$ , and define  $\sigma_{jI} = \mathbb{C}(W_{ji}, I_i)$  and  $\sigma_j^2 = \mathbb{V}(W_{ji})$ . The intuition behind the method proposed in [Altonji et al. \(2005\)](#) and extended by [Oster \(2019\)](#) is to assume that the correlation between the vectors  $W_{ji}$  and  $I_i$  is proportional—that is,

$$\delta \frac{\sigma_{1I}}{\sigma_1^2} = \frac{\sigma_{2I}}{\sigma_2^2}, \quad (6)$$

with  $\delta$  being the coefficient of proportionality. This can be rationalized by assuming a correlation structure, in a projection sense, between the vectors  $X_i^o$  and  $X_i^u$ . [Oster \(2019\)](#) assumes  $\delta = 1$ . That is to say, biases from omitting observables and unobservables are of similar magnitudes. Bounds, however, can be computed under different assumptions on  $\delta$  to accommodate different a priori correlation structures.<sup>6</sup>

With an argument based on auxiliary regressions, [Oster \(2019\)](#) characterizes OVB as a function of  $\delta$  that naturally suggests bounds on the estimated coefficient. [Oster \(2019\)](#) also notes that the degree to which the sensitivity analysis informs us about OVB depends on the variance of  $\xi_i$  or on the contribution of  $X_i$  to the variance of  $Release_i$ , conditional on  $I_i$ . Intuitively, coefficient stability may also arise from lack of explanatory power, so the implied OVB has to be scaled by the explanatory power of the observed and unobserved covariates. This implies that the econometrician also has to make an assumption about  $R_{max}$ —that is, the hypothetical  $R^2$  that would be obtained after estimating (5) by OLS including both  $X_i^o$  and  $X_i^u$ . Bias is less likely to be problematic if coefficient stability after the inclusion of  $X_i^o$  is accompanied by important increases in goodness of fit relative to  $R_{max}$ .

Table 3 presents the estimated bounds for the coefficient  $\hat{\alpha}_D$  under different assumptions on  $\delta$  and  $R_{max}$ . Each cell shows the estimated  $\hat{\alpha}_D$  parameter adjusted for OVB under the assumption

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<sup>6</sup>Assuming  $\delta = 1$  mimics a setting in which  $X_i^o$  is a random subsample of  $X_i$ . If one assumes that the most important elements of  $X_i$  are included in  $X_i^o$ , then  $\delta$  is likely to be smaller than one.

Table 3: Oster (2019) Test for Coefficient Stability:  $\hat{\alpha}_D$

$R_{max}$	$\delta$					
	-1	-0.5	-0.25	0.25	0.5	1
0.25	-0.081	-0.084	-0.085	-0.088	-0.090	-0.093
0.5	-0.056	-0.071	-0.079	-0.095	-0.103	-0.121
0.75	-0.033	-0.059	-0.072	-0.102	-0.117	-0.151
1	-0.012	-0.047	-0.066	-0.109	-0.132	-0.185

**Notes:** This table presents bounds for the estimated coefficient  $\hat{\alpha}_D$  that correct for potential OVB by implementing the method developed by Oster (2019), under different assumptions on  $\delta$  and  $R_{max}$ .  $\delta$  is the coefficient of proportionality that accounts for the correlation between observables and unobservables.  $\delta = 0$  is not included because it rules out OVB by construction.  $R_{max}$  is the  $R^2$  of a regression that includes observables and unobservables.

that selection on observables informs us about selection on unobservables. Note that larger values for  $\delta$  and  $R_{max}$  imply more conservative bounds on  $\hat{\alpha}_D$  because both impose a stronger role for unobservables in the release equation.

Imposing  $\delta > 0$ , which is the standard assumption, suggests that discrimination is underestimated: bounds are larger (in absolute value) than the fully controlled benchmark regressions displayed in Column 5 of Table 2. This follows from the fact that, on average, immigrant defendants have cleaner criminal records than Chilean defendants. It also relates to the discussion on Column 2 in Table 2, which posited that, if anything, unobservables exert downward bias on the discrimination estimates. The results show that the only way in which OVB works against finding discrimination is by assuming that  $X_i^o$  and  $X_i^u$  are negatively correlated—that is, that defendants with observables that induce lower misconduct potential also have unobservables that work in the opposite direction. Although it is counterintuitive, this scenario can be explored by assuming negative values for  $\delta$ . Note that, even in the most conservative (and presumably implausible) scenario in which the negative correlation is large ( $\delta = -1$ ) and the explanatory power of omitted variables is sizable ( $R_{max} = 1$ ), the point estimate still suggests (mild) discrimination against immigrant defendants. These results suggest that our main findings are unlikely to be driven by unobservables that explain misconduct potential and correlate with immigration status. By contrast, we provide suggestive evidence that, if anything, the results in Table 2 underestimate the level of discrimination against immigrant defendants in pretrial release decisions.

**Diagnostic 2: Using the outcome equation to test for omitted variable bias** We develop an alternative test to assess the role of unobservables in explaining the estimated release disparities. This approach is based on our theoretical framework and uses the estimation of the outcome

equation (controlling for selection bias) to validate the benchmark equation. Intuitively, relevant omitted variables predict misconduct potential and correlate with immigration status. A regression of observed pretrial misconduct against immigration status controlling for our proxies of misconduct potential should therefore inform about the partial correlation between being an immigrant and pretrial misconduct.

If we assume linear models for simplifying the exposition, then the release and outcome equations can be written as

$$Release_i = \alpha_0 + \alpha_D I_i + X_i^{o'} \alpha_{X_o} + X_i^{u'} \alpha_{X_u} + \xi_i, \quad (7)$$

$$PM_i = \beta_0 + \beta_D I_i + X_i^{o'} \beta_{X_o} + X_i^{u'} \beta_{X_u} + \epsilon_i. \quad (8)$$

In what follows, the estimated coefficients from regressions that include  $(I_i, X_i^o, X_i^u)$  are denoted by  $(\hat{\alpha}_D^{ou}, \hat{\beta}_D^{ou})$ , and the estimated coefficients from regressions that only include  $(I_i, X_i^o)$  are denoted by  $(\hat{\alpha}_D^o, \hat{\beta}_D^o)$ . Concerns about OVB in observational benchmark regressions (possibly) imply that  $\mathbb{E}[\hat{\alpha}_D^{ou}|I_i, X_i] \neq \mathbb{E}[\hat{\alpha}_D^o|I_i, X_i^o]$ . We argue that failing to reject  $\beta_D = 0$  in an outcome regression that omits  $X_i^u$  implies that  $\mathbb{E}[\hat{\alpha}_D^{ou}|I_i, X_i] = \mathbb{E}[\hat{\alpha}_D^o|I_i, X_i^o]$ ; thus, the outcome regression can be used as an indirect test for OVB in the release equation.

To see why, recall that, by assumption,  $X_i = [X_i^o \ X_i^u]$  is an accurate proxy of misconduct potential, implying that  $\beta_D = 0$ : immigration status cannot predict pretrial misconduct after controlling for misconduct potential. The hypothetical (unselected) regression of  $PM_i$  versus  $I_i$  and  $X_i$  yields unbiased estimates. Specifically,  $\mathbb{E}[\hat{\beta}_D^{ou}|I_i, X_i] = \beta_D = 0$ .<sup>7</sup> This logic does not hold in a regression that omits  $X_i^u$  because  $\mathbb{E}[\hat{\beta}_D^o|I_i, X_i^o] = \beta_D + bias$ , where the bias depends on  $\beta_{X_u}$  and the correlation between  $I_i$  and  $X_i^u$  through a standard OVB formula. Assuming  $\beta_{X_u} \neq 0$  (the contrary would imply that OVB is ruled-out by assumption), nonzero estimates of  $\hat{\beta}_D^o$  indicate that the correlation between  $I_i$  and  $X_i^u$  is nonzero. Intuitively,  $I_i$  becomes a good predictor for pretrial misconduct only if it correlates with elements of  $X_i$  that are not observed by the econometrician.

The nonzero correlation between  $I_i$  and  $X_i^u$  is exactly what induces bias when estimating discrimination in the release equation without observing  $X_i^u$ . In particular,  $\mathbb{E}[\hat{\alpha}_D^{ou}|I_i, X_i] = \mathbf{D}$  because controlling by  $X_i$  implies that  $\hat{\alpha}_D^{ou}$  identifies the direct effect of  $I_i$ , as well as the effects of other unobservables that are correlated with  $I_i$  but that are not related with misconduct potential (con-

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<sup>7</sup>This requires us to assume that  $\epsilon_i$  is uncorrelated with  $I_i$ , so  $\mathbb{E}[\epsilon_i|I_i, X_i] = \mathbb{E}[\epsilon_i|X_i] = 0$ . If we think of  $\epsilon_i$  as the projection error,  $Y_i^* = g(X_i) + \epsilon_i$ , then the second equality is true by construction, and the first equality requires us to assume that the projection error does not vary with immigration status, which is a plausible assumption given that  $X_i^u$  is unrestricted (eventually making  $\epsilon_i$  negligible) and allowed to correlate with  $I_i$ .

tained in  $\xi_i$ ) and, therefore, constitute valid sources of discrimination under a disparate impact perspective. By contrast,  $\mathbb{E}[\hat{\alpha}_D^o | I_i, X_i^o] = \mathbf{D} + \text{bias}$ , where the bias depends on  $\alpha_{X_u}$  and, again, the correlation between  $I_i$  and  $X_i^u$ . Note that the correlation that induces OVB in the release equation is the same that induces nonzero  $\hat{\beta}_D^o$  coefficients in the outcome equation.

Putting the two arguments together, we note that failing to reject  $\beta_D = 0$  in a regression that omits  $X_i^u$  (i.e., when the point estimate is  $\hat{\beta}_D^o$ ) implies  $\mathbf{D} = \mathbb{E}[\hat{\alpha}_D^{ou} | I_i, X_i] = \mathbb{E}[\hat{\alpha}_D^o | I_i, X_i^o]$ . Put simply, not rejecting that immigration status does not predict pretrial misconduct on top of  $X_i^o$  confirms that the vector  $X_i^o$  successfully controls for group differences in misconduct potential and, therefore, the discrimination estimates from equation (7) are not affected by standard OVB. One challenge in implementing this test is that the sample of released defendants is selected because pretrial misconduct if released is not observed for detained individuals. Accordingly, we estimate selection models that use the judge and attorney controls as the excluded variables in the outcome equation based on the assumption that judges' leniency and attorneys' quality affect selection into treatment but do not affect misconduct potential.<sup>8,9</sup> Our main specification implements a standard two-step parametric Heckit selection correction (Heckman, 1974). We also provide results using the semiparametric correction proposed by Newey (2009).<sup>10</sup>

Table 4 presents the results using the parametric selection correction. Table A.I of Appendix A presents the results using the semiparametric selection correction. Each column represents a different regression and reports the estimated  $\hat{\beta}_D^o$  coefficient from equation (8) with its corresponding standard error. In all regressions, the selection equation includes the complete set of controls, but the included covariates in the outcome equation vary between columns. Note that the F-test for judge leniency and attorney quality (and their squares) suggests they effectively help with the

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<sup>8</sup>To avoid saturating the nonlinear first-stage model with the court-by-time fixed effects, we replace them with time-varying variables at the court level: number of judges, average pretrial release rate, and number of prosecutions (within a court in a given year). Estimating the fully controlled benchmark regression (Column 5 in Table 2) and replacing the court-by-time fixed effects by these time-varying court-level controls yields a point estimate of -0.084, suggesting that the time-varying variables do a good job of approximating the court-by-time fixed effects.

<sup>9</sup>Although the test proposed here is different from the weighting correction made in Arnold et al. (2021), it is worth noting that their proposal is also based on the intuition that the unselected outcome equation can be used to assess the degree of OVB in the selection equation.

<sup>10</sup>The Heckit selection correction assumes that the error terms of the first and second steps are jointly normal. The semiparametric correction of Newey (2009) uses series approximations to compute control function corrections. We implement the semiparametric correction following Low and Pistaferri (2015) where the first step uses Gallant and Nychka (1987) estimator to approximate the unknown density by third degree Hermite polynomial expansions and the second step controls for non-linear transformations of the density prediction. As in Low and Pistaferri (2015), we consider three models. Let  $\hat{f}$  denote the predicted density. The control function used in Model I is  $\hat{f}$  and its square, in Model II is  $\Phi(\hat{\alpha}_0 + \hat{\alpha}_1 \hat{f})$  and its square—where  $\Phi$  is the normal cumulative distribution function and  $(\hat{\alpha}_0, \hat{\alpha}_1)$  are the estimated coefficients of a Probit model of *Release* on a constant and  $\hat{f}$ —, and in Model III is  $\lambda(\hat{\alpha}_0 + \hat{\alpha}_1 \hat{f})$  and its square—where  $\lambda(x) = \phi(x)/\Phi(x)$  is the inverse Mills ratio and  $\phi$  the normal density.

Table 4: Parametric Selection Model for Assessing OVB:  $\widehat{\beta}_D^o$

	(1)	(2)	(3)	(4)
Immigrant	-0.078 (0.007)	-0.012 (0.006)	-0.090 (0.007)	-0.008 (0.006)
Mean dep. variable	0.29	0.29	0.29	0.29
Individual controls	No	Yes	No	Yes
Court-by-year controls	No	No	Yes	Yes
No. of Immigrants	4,900	4,900	4,900	4,900
No. of Chileans	580,406	580,406	580,406	580,406
F-test for excluded variables	969.6	969.6	969.6	969.6

**Notes:** This table presents the results from the two-step parametric sample selection model (Heckman, 1974) with different sets of controls, where the dependent variable is pretrial misconduct. We report the point estimate for the immigrant indicator (i.e., the coefficient  $\widehat{\beta}_D^o$ ) of equation (8) and its standard error. Both sets of controls (*individual controls* and *court-by-year controls*) are always included in the selection equation, but the columns vary in their inclusion in the outcome equation. Judge and attorney controls are defined as in Table 2 and are excluded from the outcome equation. The last row presents the joint F-test for judge and attorney (excluded) controls in the selection equation. *Individual controls* are defined as in Table 2. To avoid saturating the nonlinear first-stage with court-by-year fixed effects they are replaced in the regressions by court-by-year time varying covariates—namely, the average number of judges, the average pretrial release rate, and the number of prosecutions (within a court in a given year).

selection correction.

Column 1 in Table 4 shows that, after controlling for selection, pretrial misconduct rates are 7.8 pp smaller for immigrant defendants. This suggests that the benchmark regression with no controls (Column 1 in Table 2) is affected by OVB. The implied OVB, however, suggests that abstracting from controls leads to an underestimation of the discrimination against immigrant defendants. Column 2 shows that including individual controls (criminal records and type of crime) reduces the pretrial misconduct rate differences to 1.2 pp. That is to say, conditioning on the individual controls reduces group variation in  $Y_i^*$  by 85%. This suggests that, although imperfect, criminal records and type of crime are good predictors of pretrial misconduct potential. Adding time-varying court controls do little to correct for unobserved differences in misconduct potential (Column 3); however, adding both individual and time-varying court controls reduces the misconduct rate disparities to a nonsignificant gap of less than a percentage point. Taken together, we interpret this as support for our choice of  $X_i^o$ . The vector  $X_i^o$  seems to do a good job in accounting for group differences in misconduct potential, which implies that the discrimination estimates of Column 5 of Table 2 are not driven by unobserved variation in latent risk. The model with the semiparametric selection correction proposed by Newey (2009) yield the same conclusions.<sup>11</sup>

<sup>11</sup>The three models reproduce the pattern displayed in Table 4. After accounting for selection, immigrants have lower pretrial misconduct rates. These gaps are substantially reduced after including individual controls in the

A similar test can be implemented by comparing the role of unobservables in the selection and outcome equations using Kitagawa–Oaxaca–Blinder (KOB) decompositions (Kitagawa, 1955; Oaxaca, 1973; Blinder, 1973). Intuitively, unobservables can affect the interpretation of the estimated benchmark coefficients if they correlate with immigration status and explain misconduct potential. If unobservables are important for predicting group-specific misconduct potential when making release decisions, they should also explain differences in observed pretrial misconduct rates.<sup>12</sup> Appendix C shows KOB decompositions based on the vectors of individual covariates, using both raw data and residualized data against court-by-time fixed effects. The share of variation that is not explained by observables is large in the release equation but it is essentially zero in the outcome equation. This suggests that the immigrant indicator in the release equation is effectively capturing discrimination and not differences in unobservables.

**On included variable bias** Our diagnostics do not explicitly deal with IVB concerns, but the second OVB diagnostic provides some insights to assess the pervasiveness of IVB. Individual controls are much more important than court-by-time variation in terms of accounting for group differences in pretrial misconduct rates. This indicates that the inclusion of court-by-time fixed effects in the benchmark regression may induce IVB in the sense that these particular fixed effects impact the estimated discrimination coefficient but do little to attenuate OVB. In other words, regressions that include court-by-year fixed effects could be over controlling and therefore inducing bias to our discrimination estimates.

IVB affects the interpretation of the sensitivity analysis. As shown in Table 2, including court-by-year fixed effects significantly reduces the discrimination estimate. The results presented in Table 4 suggest that the sensitivity of  $\hat{\alpha}_D$  to the inclusion of court-by-year fixed effects is not explained by unobserved variation in misconduct potential but rather by an alternative mechanism that mediates discrimination against immigrant defendants. For example, the fact that discrimination decreases when controlling by court-by-time fixed effects may indicate that part of the discrimination against immigrant defendants is driven by their systematic allocation to courts that are more severe on

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outcome equation. In Model I, individual controls reduce the gap by two thirds, although the gap remains significant at 3 pp after including the complete set of controls. In Models II and III the individual controls completely close the gap in pretrial misconduct to a nonsignificant estimate of less than a percentage point, even in the specification that excludes time-varying court characteristics.

<sup>12</sup>One caveat of this analysis is that the sample used in the outcome equation is selected, presumably based on unobservables. To the extent that the selected sample has variation in unobservables, however, the exercise should still capture their effects. This would not be the case if, for example, the release decision is based on a binary unobservable such that all released defendants have the same value. In this case, the lack of variation in the selected sample would prevent the decomposition from capturing the effect of unobservables—that is, the KOB exercise could be noninformative under specific patterns of selection on unobservables.

average. Therefore, from a disparate impact perspective, the relevant discrimination estimate could be the one that only includes individual controls (Column 2 of Table 2), which suggests larger levels of discrimination against immigrant defendants than the fully controlled specification. Because this conjectured IVB attenuates the discrimination estimate, we take a conservative position by presenting the fully controlled benchmark regression as our preferred specification. It is important to keep in mind, however, that the impact of IVB may mean that the actual discrimination against immigrant defendants is even larger.

## 5 Dissecting the Discrimination Estimates

Section 4 documents a robust pretrial release rate disparity between immigrant defendants and Chilean defendants after controlling for variables that proxy for misconduct potential that we interpret as evidence of discrimination against immigrant defendants. This section dissects this reduced-form result to get a more comprehensive characterization of the discrimination patterns against immigrant defendants. All results presented in this section consider the full set of controls and should therefore be compared to Column 5 in Table 5.

First, we explore some behavioral drivers of the discrimination estimate. We show that a sizable portion of the effect stems from an informational problem: immigrant criminal records are censored and this leads judges to punish immigrants with no previous prosecutions. Outcome test estimations show that the result cannot be rationalized by accurate statistical discrimination, suggesting that biased beliefs and taste-based discrimination play a role in the resolution of the informational problem. Second, we explore heterogeneities by type of crime and find that discrimination is especially large for drug-related offenses. This is consistent with a hostile discourse in the public sphere against immigrants that links immigration to drug trafficking. Finally, we explore the time trend of discrimination patterns considering the recent immigration wave discussed in Section 2. We find that discrimination has dramatically increased in recent years, and that this increase is not explained by the change in the composition of the immigrant population.

### 5.1 Behavioral Drivers of Discrimination

The criminal records of immigrant defendants are censored (i.e., judges cannot observe the criminal histories of immigrant defendants in their origin countries). This means that judges may interpret the information available to them differently. In Bayesian jargon, the quality of the signal may differ between Chilean defendants and immigrant defendants. This is likely to be more detrimental

Table 5: Behavioral Drivers of Discrimination and the Role of Information

	Benchmark Test			Outcome Test	
	Previous Prosecution			KPT (4)	GV (5)
	Yes (1)	No (2)	(3)		
Immigrant	-0.019 (0.007)	-0.132 (0.015)	-0.135 (0.023)	-0.050 (0.007)	-0.074 (0.016)
Immigrant x experienced judge			0.049 (0.022)		
Mean dep. variable	0.80	0.92	0.82	0.29	0.37
No. of Immigrants	2,630	3,732	4,187	4,900	840
No. of Chileans	472,636	220,370	357,520	580,406	57,690
R-squared	0.18	0.17	0.21		

**Notes:** Columns 1 to 3 present the estimated benchmark coefficient,  $\hat{\alpha}_D$ , of the extensions of equation (5) and its corresponding standard error (clustered at the court-by-year level). Each column represents a different regression and includes all sets of controls analogous to Column 5 in Table 2 (see Table 2 notes for details). Columns 1 and 2 present  $\hat{\alpha}_D$  separately for defendants with and for defendants without previous prosecutions. Column 3 presents  $\hat{\alpha}_D$  for the period 2013–2017 and includes an interaction between  $I_i$  and an indicator that takes the value one if the case was handled by an experienced judge (i.e., judges that handled 200 or more cases in the 2008–2012 period). Columns 4 and 5 present the outcome test estimations using Knowles et al. (2001) (KPT) and Grau and Vergara (2021) (GV) approaches, respectively. KPT approach consists on an OLS regression of pretrial misconduct on the immigrant indicator for the complete sample of released defendants. GV performs a sample selection procedure that limits the extent of inframarginality bias in the KPT implementation. Details of the outcome test implementation of GV can be found in Appendix D.

to immigrant defendants with no criminal history in Chile because the main information contained in the criminal records is at the extensive margin. We explore whether this informational gap is in fact driving discrimination patterns against immigrant defendants.

To test the information hypothesis, we estimate the main benchmark regressions separately for defendants with and for defendants without previous prosecutions. These estimates include all sets of controls analogous to Column 5 in Table 2. Results are displayed in Columns 1 and 2 in Table 5. Immigrant defendants with previous prosecutions are 1.9 pp less likely to be released compared to Chilean defendants with previous prosecutions; compared to a difference of 13.2 pp for defendants without previous prosecutions. To assess the magnitude of the difference between the estimations, note that the (unconditional) pretrial release rate is 12 pp larger for defendants with clean criminal records, which is roughly similar in magnitude to the difference in the discrimination estimate between samples. This implies that not having previous prosecutions is, essentially, irrelevant for immigrant defendants: they are not “rewarded” for having clean criminal records. This result is consistent with judges systematically imputing nonzero criminal histories for all immigrant defendants regardless of their criminal record in Chile.

To further explore this intuition, we test whether judge experience matters for the aggregate discrimination patterns. If judges face an informational problem whose imperfect solution leads to discriminatory practices, they should gradually implement nondiscriminatory solutions as they gain experience. Consequently, we should observe that more experienced judges exhibit lower levels of discrimination. To test this, we define experienced judges as those who handled more than 200 cases in the period 2008–2012 and estimate the benchmark regressions for the period 2013–2017, including the experience indicator and its interaction with the immigrant indicator. Column 3 shows that release disparities are around 35% smaller for experienced judges.

As discussed in Section 3, aggregate patterns of discrimination represent a possibly complex combination of different behavioral sources—these being, statistical discrimination, biased beliefs, and taste-based discrimination, which is not possible to identify separately without additional structural assumptions (Arnold et al., 2021; Bohren et al., 2021; Hull, 2021). The fact that discrimination in pretrial detention decisions stems from an informational problem does not necessarily mean that the aggregate estimate can be rationalized by accurate statistical discrimination. On the one hand, judges can circumvent the informational problem based on stereotypes. This may, for example, lead judges to “overcorrect” for censored criminal records for immigrants. On the other hand, the informational problem may provide judges with the room to exercise taste-based discrimination. The fact that experienced judges still display significant discrimination against immigrants suggests that these drivers are likely to play a role in the analysis.

To get additional insights, we implement the outcome test (Becker, 1957, 1993) using the observational approaches of Knowles et al. (2001) and Grau and Vergara (2021).<sup>13</sup> The outcome test is a diagnostic for differences in effective selection thresholds between groups. Using the notation set out in Section 3 (equation (2)), the test explores whether  $\mathbb{E}[t_{j(i)}(1, Z_i) - b_{j(i)}(1, Z_i)]$  is different from  $\mathbb{E}[t_{j(i)}(0, Z_i) - b_{j(i)}(0, Z_i)]$ , where the expectation is taken across  $Z_i$  and  $j(i)$ . Differences in effective thresholds can be identified by differences in observed misconduct among marginally released defendants (Arnold et al., 2018; Hull, 2021). Knowles et al. (2001) propose using the average behavior of released defendants. Grau and Vergara (2021) refine Knowles et al. (2001) approach by performing a sample selection procedure that limits the extent of inframarginality bias.<sup>14</sup>

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<sup>13</sup>The small share of the immigrant population prevents us from implementing the instrument-based approach proposed by Arnold et al. (2018).

<sup>14</sup>When risk distributions vary by group, the behavior of the average released defendant is potentially different from the behavior of the marginally released defendants. This is called inframarginality bias. Grau and Vergara (2021) deal with this problem in two stages. First, it uses observational variation to identify samples of marginal defendants. The identification argument works under sufficient conditions which are supported by suggestive evidence presented in Appendix D. Second, it performs differences in means of pretrial misconduct rates between immigrant defendants and Chilean defendants using the samples of marginal defendants. The sample of marginal defendants is defined as

Intuitively, discrimination in pretrial detention decisions may originate from judges setting stricter release thresholds for immigrant defendants, in a disparate impact sense, that imply lower release rates at equivalent true pretrial misconduct probabilities. These differences in the treatment of immigrant defendants could be driven by stereotypes or animus or both. This source of discrimination does not incorporate sources of release disparities driven by differences in group-specific misconduct potential. Differences in true pretrial misconduct potential affect how often a given defendant crosses the release threshold, but the threshold itself is not informative about them. Then, rejecting the outcome test implies that the observed behavior of marginally released defendants is inconsistent with a selection process that lacks stereotypes and animus (Hull, 2021).

Columns 4 and 5 in Table 5 show the results of the outcome test using Knowles et al. (2001) and Grau and Vergara (2021) approaches, respectively. Estimates indicate that immigrant defendants face effective release thresholds that are, on average, between 5 and 7.2 pp stricter. That is to say, immigrants are required to have, on average, smaller pretrial misconduct probabilities to qualify for pretrial release relative to Chilean defendants. Although the magnitudes between benchmark and outcome regressions are not directly comparable (Domínguez et al., 2022), this result rejects accurate statistical discrimination as the only driver of the discrimination estimates, implying that a combination of stereotypes and animus plays an important role in explaining aggregate patterns of discrimination against immigrants.<sup>15</sup>

## 5.2 Heterogeneity by Type of Crime

If statistical structures, stereotypes, or judges preferences vary with type of crime, then discrimination patterns should do as well. Exploring this particular pattern of heterogeneity is instructive because a hostile facet of the public discourse regarding immigrants links immigration with drug trafficking and violent crimes, so discrimination should be particularly salient in pretrial detention decisions concerning these crimes.

To investigate this, we split prosecutions between the following nine mutually exclusive categories: homicides, sexual offenses, thefts and robberies, other property crimes, drug offenses, white collar crimes and tax crimes, crimes against public trust, crimes against the freedom and privacy

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the 10% of released defendants with larger conditional probabilities of being marginal. That is why the number of observations is smaller when using this approach. Additional details about the implementation of the outcome test proposed by Grau and Vergara (2021) are presented in Appendix D.

<sup>15</sup>The average pretrial misconduct rate is larger in Column 5 (37%) relative to Column 4 (29%). This suggests that Grau and Vergara (2021) implementation of the outcome test successfully attenuates inframarginality concerns by focusing on a sample of released defendants that are, on average, riskier and, therefore, presumably more likely to be at the margin of pretrial detention.

Table 6: Discrimination by Type of Crime:  $\hat{\alpha}_D^k$

	Coeff.	SE
Homicide	-0.030	(0.067)
Sexual offense	0.001	(0.033)
Theft or robbery	-0.038	(0.018)
Other property crime	-0.052	(0.014)
Drug offense	-0.351	(0.034)
White collar or tax crime	-0.020	(0.019)
Crime against public trust	0.021	(0.015)
Crime against people’s freedom and privacy	-0.007	(0.006)
Other crimes	0.019	(0.021)
Mean dep. variable	0.84	
No. of Immigrants	6,362	
No. of Chileans	693,006	
R-squared	0.192	

**Notes:** This table presents the estimated benchmark coefficients by type of crime,  $\hat{\alpha}_D^k$ , of equation (9) and its corresponding standard error in parentheses (clustered at the court-by-year level). Each column represents a different regression and includes all sets of controls analogous to Column 5 in Table 2 (see Table 2 notes for details).

of people, and other crimes.<sup>16</sup> Using these categories, we estimate the following regression:

$$Release_i = \alpha_0 + \sum_{k=1}^9 \alpha_D^k I_i \cdot C_i^k + X_i^{o'} \alpha_{X_o} + \varepsilon_i, \quad (9)$$

where  $C_i^k$  is an indicator that takes the value one if the imputed crime is in category  $k$ . All other variables are defined as in equation (5); therefore, the noninteracted crime indicators are included in  $X_i^o$ . In this specification,  $\hat{\alpha}_D^k$  identifies the controlled pretrial release disparities between Chilean defendants and immigrant defendants (our measure of discrimination) in crime category  $k$ .

Table 6 shows the results. Discrimination against immigrants is especially large for drug offenses—around four times larger than the baseline benchmark estimate. Significant effects are also found for thefts and robberies and property crimes. Other violent crimes like homicides and sexual offenses show no significant differences; however, as shown in Table A.II of Appendix A, the number of immigrants imputed for sexual offences and homicides in the period considered is too low, therefore potentially being underpowered to precisely estimate disparities in those particular categories. Crime categories that are associated with less hostile discourses in the public sphere (such as white collar crimes and tax crimes or crimes against privacy, and for which the number of

<sup>16</sup>For examples of the crimes included in each category, see Appendix B.

Table 7: Discrimination in Different Periods and Group Heterogeneity:  $\hat{\alpha}_D$ 

	2008-2012		2013-2017	
	(1)	(2)	(3)	(4)
Immigrant	-0.066 (0.016)		-0.101 (0.015)	
Immigrant (old group)		-0.064 (0.017)		-0.099 (0.015)
Immigrant (new group)				-0.108 (0.022)
Mean dep. variable	0.85	0.85	0.82	0.82
No. of Immigrants (old group)	1,956	1,956	2,873	2,873
No. of Immigrants (new group)	219	0	1,314	1,314
No. of Chileans	335,486	335,486	357,520	357,520
R-squared	0.172	0.172	0.207	0.207

**Notes:** This table presents the estimated benchmark coefficient,  $\hat{\alpha}_D$ , of extensions for equation (5) and its corresponding standard error (clustered at the court-by-year level). Each column represents a different regression and includes all sets of controls analogous to Column 5 in Table 2 (see Table 2 notes for details). Columns 1 and 2 use data from the period 2008–2012, and Columns 3 and 4 use data from the period 2013–2017. In columns 2 and 4 the immigrant indicator is defined separately depending on the nationality of the immigrant. *New group* includes immigrants from Colombia, Haiti, and Venezuela. *Old group* considers immigrants from all the other countries.

prosecuted immigrants is in fact larger) also show no significant release disparities. Overall, these results suggest that discrimination patterns are mainly driven by drug offenses, property crimes, and thefts and robberies. This is consistent with idea that public discourses regarding immigrants play a role in the formation of stereotypes and hostile preferences.

### 5.3 Time Trend and the Immigration Wave

Recall in Section 2 it was premised that immigration flows have substantially increased in recent years, generating a hostile discourse against immigrants in the public sphere. It was also argued that the composition of the immigrant population by country of origin has changed considerably. We now examine whether these changes have been accompanied by a change in the aggregate discrimination patterns across time.

Following the changes in immigration flows described in Section 2, Table 7 shows the benchmark estimates separately for the periods 2008–2012 and 2013–2017. The fully controlled discrimination estimate increased from 6.6 to 10.1 pp, a 53% increase. A potential explanation for this increase is the increase in the hostility of the public discourse that may have affected stereotypes and preferences. Alternatively, this result could be driven by a composition effect. Although Peru

was the principal source of immigration flows to Chile for many years, beginning in 2013 the relative share of immigrants from Colombia, Haiti, and Venezuela increased considerably. If there is heterogeneity in discrimination patterns—that is, discrimination is more intense against the “new” group of immigrants (for example, because of more salient racial differences)—then the increase of discrimination between the two periods could be explained by a composition effect and not by an increase in the average intensity of discrimination for a given nationality.<sup>17</sup> Column 4 presents the discrimination estimate for the 2013–2017 period separately for the “new” and the “old” groups of immigrants.<sup>18</sup> We cannot reject that the discrimination estimates for both groups are equal, which suggests that the increase in discrimination is unlikely to be driven by the change in the composition of the immigrant population.

## 6 Conclusion

This paper leverages rich administrative data to test for discrimination in pretrial detention decisions against defendants who are immigrants in Chile. We find that immigrant defendants are released pretrial at substantially lower rates relative to Chilean defendants with similar criminal records, case characteristics, and court-by-time fixed effects. Several diagnostics suggest that these estimated disparities are not driven by unobserved differences in pretrial misconduct potential between Chilean defendants and immigrant defendants. We therefore interpret the results as robust evidence of discrimination against immigrant defendants in pretrial detention decisions.

We show that release disparities are an order of magnitude larger for defendants with no previous criminal prosecutions in Chile. We argue this implies that discrimination is originated in an informational problem driven by censored criminal records (or, put differently, weaker signal quality). However, we show that accurate statistical discrimination cannot rationalize the release disparities, suggesting that biased beliefs and taste-based discrimination play an important role. Consistent with the hostile discourses against immigrants in the public sphere, we also find that discrimination in pretrial detention decisions is particularly severe for drug offenses, and that the intensity of discrimination has increased in recent years during which the magnitude and the composition of the immigrant population has changed considerably.

Overall, our findings suggest that discrimination against immigrant defendants in pretrial de-

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<sup>17</sup>Potential differences in risk between the two groups are not a concern for the dynamic results because benchmark regressions include the full set of controls.

<sup>18</sup>This exercise cannot be computed for the 2008–2012 period because the “new” group, by definition, is essentially nonexistent before 2013, as the number of observations suggest.

tention decisions is substantial in the Chilean context and highlight the importance of implementing effective anti-discrimination policies to better assimilate the integration of immigrants in destination countries. Several countries are enacting anti-discrimination policies (or strengthening existing ones) to foster the integration of immigrant populations (OECD, 2019). This policy challenge appears to be particularly urgent in countries where immigration is a relatively recent phenomenon and institutions are not adequately equipped to deal with unexpected shocks in this regard.

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