Prices and Policies in Opioid Markets*

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Abstract

Choice between prescription and illicitly-manufactured sources of opioids for nonmedical use is modeled, with emphasis on the potential nonconvexity of consumer budget sets. The theory helps reconcile apparently disparate empirical findings, identify groups whose price responses are opposite of the average, and draw lessons from previous studies even for policy environments with new supply conditions. Organized around the model's two supply channels, an opioid policy database is assembled that helps reveal distinct pricing phases during the years 1999-2021. Consistent with the model, during the later pricing phases the cross-area relationship between the opioid fatality rate (measured with death-certificate data from the Centers for Disease Control and Prevention) and its composition changes sign; fatality rates among children and youth trend opposite of adult rates; and the black-white fatality gap changes sign, especially among older people.

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I. Introduction

In both 2015 and 2016, U.S. life expectancy fell from the previous year. A single year drop had not happened in 22 years, and two consecutive declines had not occurred in more than 50 years. The sharp reversal in the national trend toward longer lives is widely understood to be connected to the opioid epidemic, whose annual U.S. costs are approaching a trillion dollars. A similar reversal may be soon observed in other countries and regions where fatalities involving opioids have already increased by several multiples in a decade or so.[1] The fatalities likely indicate millions more consumers who still struggle with opioid addictions.

This paper presents an economic model of choosing between prescription (Rx) and illicitly-manufactured (Im) sources of opioids for nonmedical use.[2] Although simple, it unifies and helps explain a range of policy effects that have been documented in the literature, as well as new empirical results. The model also shows what previous findings on, say, prescription regulation, may reveal about other technological and regulatory changes in opioid markets that would appear unrelated to prescriptions. The predictions of an economic model are especially valuable for opioid markets where data can be sparse and policy analysis might rationally put more weight on potentially relevant lessons from other contexts.

Medical experts advising or serving as policymakers typically ignore the interplay between Rx and Im delivery channels. As recently as 2022, the Stanford-Lancet Commission on the North American Opioid Crisis recommended changes in law enforcement and prescription regulations without acknowledging that their proposals might increase both demand and supply in illicit markets (Humphreys, et al. 2022). The U.S. Food and Drug Administration, which

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[1] Opioid death rates increased by a factor of about six in Sweden, Northern Ireland, and British Columbia, surpassing by 2018 or 2019 the rates that the U.S. had as recently as 2013 (Pardo 2019, Chapter 4, Northern Ireland Statistics and Research Agency 2020). Period life expectancy is FRED series SPDYNLE00INUSA. Opioid costs are from Murphy (2020), which include value of lost lives and other costs but no offset for "consumer surplus."

[2] Opioids include prescription painkillers such as oxycodone (an active ingredient in Oxycontin and Percocet) and hydrocodone (an active ingredient in Vicodin) as well as morphine and illicitly-manufactured drugs such as heroin, illicit fentanyl, and fentanyl analogs.

oversees the marketing of prescription products, refuses to consider any costs that accrue in heroin or fentanyl markets, because those markets are outside their jurisdiction (Mulligan 2020).

The economic model suggests a different approach. With multiple sources of opioids, consumers potentially face a nonconvex budget set, with a high marginal price at low levels of opioid consumption and a low price at high levels. The opioid source and the quantity consumed are simultaneously determined. A change in either price has two consequences for market aggregates: a jump from one part of the budget set to another among consumers who were indifferent between the Rx and Im sources and the ordinary movement along individual-level demand curves among those who were not indifferent. The former is a large change among a few consumers while the latter is a relatively small change among many. The two can be in opposite directions, and either can dominate in the aggregate. A contribution of this paper is a sufficient-statistics expression for comparing the two magnitudes, in both the short and long run, and for identifying groups for whom one or the other effect is especially likely.

A previous econometric literature has already warned that policies aimed at *reducing* prescription opioid consumption can lead to *increased* mortality in the short run due to widespread substitution to illicit opioids. Many of the papers provide convincing evidence that this may be the case in the United States in 2011 and subsequent years, often citing "existence" or "availability" of heroin as a critical factor driving this result.[3] At the same time, increased mortality in an earlier era, when heroin was also available, has been attributed to just the opposite: policies and business practices that increased prescription consumption.[4] The economic model clarifies that switching sources is not merely regulatory avoidance but also changes the quantity consumed among those who switch. It therefore points to the price gap between Rx and Im opioids as the critical determinant of both the sign and magnitude of the effects of prescription policies.[5] Although measuring illicit prices is subject to significant

[3] See Powell and Pacula (2021), Powel, Alpert, Pacula (2019), Meinhoffer (2018) and the studies cited in Section IV. Maclean et al (2020) is a survey of opioid economics generally.
[4] Pacula and Powell (2018) and Alpert et al. (2022).
[5] Evans, Lieber, and Power (2019) and others discuss an independent influence of heroin prices on drug consumption and mortality but not on the sign or magnitude of the effect of Rx regulation. No analysis of the gap between heroin and Rx prices is attempted.

measurement error, it is generally understood that that prescription opioids were once "poor man's heroin" but more recently "heroin is cheaper and easier to get than prescription opioids".[6]

More important, the model adds valuable predictions. It shows which groups may experience reduced mortality from prescription regulation even if the general population does not. It shows how much the current Im market must change to bring opioid markets back to a more conventional era when opioid mortality varies inversely with Rx-opioid prices. It even tightly links the consequences of Rx-price changes such as those associated with regulation; effects of Im-price changes resulting from technological progress in illicit-opioid manufacturing or the treatment of health conditions resulting from intravenous drug use, or from changes in law enforcement; and effects of price changes common to the two sources including opioid-overdose treatments and changes in labor market opportunity costs of opioid consumption. When only one set of consequences is part of the available evidence base, the link strengthens policymaking that would otherwise be limited to post-mortem analysis: waiting for mortality to accumulate before reaching conclusions about the other sets (Ruhm 2019a). A conceptual framework that sheds light on the generalizability of the historical evidence base helps save lives, especially in a market where even the direction of policy effects varies over time. To use a metaphor, confusing the policy gas and brake pedals is a tragic mistake that a conceptual framework helps avoid.

With the exception of Schnell (2018), the economics literature on prescription regulation had no formal analysis of opioid consumption incentives.[7] "Elasticities" are sometimes part of the discussion, but no indication is given as to which demand or supply elasticities are needed to explain why opioid markets reached the point that opioid mortality would increase with Rx prices. An economic model helps reconcile apparently disparate findings and draw lessons from previous studies even for policy environments with new supply conditions.

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[6] U.S. Department of Justice, National Drug Intelligence Center (2001) and National Institute on Drug Abuse (2018), respectively.

[7] Schnell (2018) builds an equilibrium model of switching between primary Rx and (unlawful) secondary Rx markets that shows conditions under which the two would be close substitutes. Greenwood et al (2022) is a new paper modeling transitions between medical and non-medical prescription use. CEA (2019, Figure 10) is a demand-residual analysis quantifying the degree to which declining Rx prices explain rising Rx mortality. The formal analysis in the rest of the literature is so far confined to equations showing econometric specifications (data construction, the use of fixed effects, etc). There are several formal models of drug demand generally.

Section II's model offers six empirical predictions in the form of formal propositions. The death-certificate and opioid-price data are described in Section III, including a new federal policy database that helps reveal distinct pricing phases during the period 1999-2021. The four propositions that are testable with that data are the subject of Section IV. One test is that that cross-area relationship between the opioid fatality rate and its composition changed at the same time that Im prices fell and Rx regulations were tightened. A second test is whether opioid deaths fell coincident with the OxyContin reformulation for children and youth, whose opioid consumption appeared to be especially Rx intensive. Both predictions are confirmed in the data. The second empirical finding at least directionally supports the gateway hypothesis that Rx regulation reduces the rate of opioid initiation and thereby might sufficiently reduce future opioid demand to have a net negative mortality effect in the long-run.

By 2008, African-Americans stood out as having lower opioid mortality rates relative to whites. If much of the differential was due to differential access to prescriptions, then the third model prediction is that black mortality rates would eventually surpass white rates once Im prices fell enough. A fourth prediction assesses the pace at which the race-reversal would occur. Section IV finds that black mortality rates did in fact surpass white rates, holding constant gender, age group, and geographic area. Moreover, as predicted, the race reversal occurred first among older people and at roughly the predicted pace. I also find that part of the race reversal at the national level is explained by the differing black-population shares across areas. Section V discusses possible model extensions, followed by the concluding Section VI.

II. Opioid Policies and the Consumer Budget Set

Model agents have preferences over two composite commodities: opioids Q and "all other goods" z. The preferences are represented by the function $u(Q,z;\theta)$, where u is strictly quasiconcave in Q and z. The scalar θ is a shifter of the marginal rate of substitution used for some of the derivations as well as representing the influence of opioid consumption in the past, as it does in models of habit, addiction, and drug tolerance (Becker and Murphy 1988, Pollak 1970). The rate of exchange between the composites is the full price of opioids, including consumer time, effort, and stigma as well as out-of-pocket costs. Although the nonlinearity of the budget constraint is essential, the indirect utility function $v(p_Q, y) \equiv \max_{Q \geq 0} u(Q, y - p_Q Q)$ for a hypothetical consumer with preferences $u(Q,z;1)$ and facing a linear budget constraint $y = z + p_Q Q$ – illuminates the derivations by summarizing the relevant features of u.

Household production.—I distinguish two broad categories of opioids: prescriptions (Rx, including prescriptions diverted into secondary markets or passed through social networks) versus illicitly manufactured (Im, especially heroin and fentanyl). On the household production side, Q is produced according to a homogeneous function $Q(q_R,q_I)$ of the Rx and Im quantities, respectively, with (at least) the Rx quantities measured in morphine-gram equivalents (MGEs). I normalize the units of Q as $Q(1,0) = 1$, so that Q's units can also be interpreted as MGEs. Finally, the units of Im are normalized so that $Q(0,1) = 1$, which means that the scale of Im measurement is proportional to MGEs but the proportionality factor may differ from one. Each MGE of Im opioids may be more productive in preferences than a MGE of Rx opioids due to the fact that Im opioids are more typically delivered intravenously. On the other hand, consumers may prefer a MGE of Rx to a MGE of Im because Im products may be less uniform and less reliable in terms of their potency and use of additives (Galenianos & Gavazza, 2017). A habit of intravenous delivery is also associated with various health problems.[8]

The uniformity, reliability, delivery, and other properties of Rx and Im are also reasons why my specification $Q(q_R,q_I)$ allows for the possibility that the two are imperfect substitutes in preferences. The elasticity of factor substitution in Q is not necessarily constant, but it exceeds one (so that purchasing just one of the two is optimal in some circumstances) and exceeds the elasticity of substitution in u. In other words, I assume that Im is a better substitute for Rx than it is for other goods. A special case of this framework has the function Q as the simple sum of the two quantities, which may be especially relevant for the high-volume consumers whose preferences heavily emphasize morphine-like symptoms over all other goods, consequences, etc.

Each of the quantities (q_R, q_I) has its own fixed cost (f_R, f_I) and marginal price (p_R, p_I), respectively. The marginal prices, which quantify the amount of other goods that are foregone by consuming one more unit of the corresponding opioids, are always positive. Particularly relevant for opioid markets is the difference $f_I - f_R$, which I expect is often (but not always)

[8] They include serious and sometimes fatal infections such as HIV, hepatitis C virus, and necrotizing soft tissue infections (Collier, Doshani and Asher 2018, Powell, Alpert and Pacula 2019, May, et al. 2021, Hrycko, et al. 2022).

positive due to Im costs of avoiding theft, acquiring self-dosing skills, or overcoming fear of needles. Moreover, because illicit-market prices are typically high and quality low for first-time buyers (Galenianos & Gavazza, 2017), establishing a trusting relationship with a drug dealer can itself be a fixed cost necessary to access a low quality-adjusted price. Depending on market conditions, the marginal price per morphine equivalent may be less for Im than Rx opioids.[9]

The consumer's choice problem is the allocation of his income among expenditures on other goods z and the fixed and variable costs of obtaining opioids, given $p_R > 0$, $p_I > 0$, $f_R \geq 0$, $f_I \geq 0$ and $y > 0$, as described in (1).

$$\begin{gathered}\max_{q_R, q_I, \phi, z} u(Q(q_R, q_I), z; \theta) \;\; s.t. \\ z \geq 0, q_R \geq 0, q_I \geq 0, \\ z + p_R q_R + p_I q_I + \phi \leq y, \\ \phi = 0 \;\; if \;\; q_R = 0 = q_I, \\ \phi = f_R \;\; if \;\; q_R > 0 = q_I, \\ \phi = f_I \;\; if \;\; q_R = 0 < q_I, \\ \phi = f_R + f_I \;\; if \;\; q_R > 0 \wedge q_I > 0\end{gathered} \tag{1}$$

where ϕ denotes the fixed costs, if any, that the consumer chooses to pay.

LEMMA 1 (Piecewise linear budget set). Let $C(Q,p_R,p_I;f_R,f_I)$ denote the minimum expenditure $p_R q_R + p_I q_I + \phi$ required to achieve output $Q \geq 0$ given p_R, p_I, f_R, f_I, and constrained by $Q(q_R,q_I) \geq Q$, $q_R \geq 0$, $q_I \geq 0$, and the four possibilities for ϕ listed in (1). Then

a. $C(Q,p_R,p_I;0,0) = Q\, C(1,p_R,p_I;0,0) \leq Q \min\{p_R,p_I\}$.
b. The consumer's budget set is $z + C(Q,p_R,p_I;f_R,f_I) \leq y$, $Q \geq 0$, and $z \geq 0$. Its boundary is piecewise linear in first quadrant of the $[Q,z]$ plane, formed as the upper envelope of the three linear budget constraints corresponding to the three fixed-cost decisions: $y = z + f_R + Q\, p_R$, $y = z + f_I + Q\, p_I$, and $y = z + f_R + f_I + Q\, C(1,p_R,p_I;0,0)$, respectively.

Proof. Because household production is homogeneous, either: one of the cost-minimizing quantities is zero and the other equal to Q, or the cost-minimizing ratio q_R/q_I is

[9] Im marginal prices can be low because, for example, the Im sector does not pay taxes and spends little on packaging. The typical delivery of Im opioids is intravenous, which may itself lower the marginal price of morphine symptoms. On the other hand, illegal sellers may forgo economies of scale in order to avoid detection by law enforcement (Campana 2016).

strictly positive regardless of Q. Either possibility yields the equality in (a). The weak inequality must hold because setting either q_R or q_I to zero is in the feasible set. An allocation $\{Q,z\}$ satisfies the set described (b) iff it is part of an allocation satisfying (1)'s constraints because C satisfies by construction. A piecewise linear boundary follows from (a). QED

Lemma 1b says that the solution to (1) can be described in two stages. In one stage, the consumer decides how to produce Q from Rx and Im, which is the minimization that defines C. In the other stage, the consumer allocates his income y between opioids Q and all other goods according to his preferences $u(Q,z;\theta)$ subject to the constraint that $z + C(Q,p_R,p_I;f_R,f_I)$ does not exceed his income. Part of the consumer's minimization problem embedded in the cost function C is whether to pay the fixed cost for Rx, the fixed cost for Im, or both in which cases he would consume only Rx, only Im, or both, respectively. The marginal cost of opioids, $\partial C/\partial Q$, is constant in each of these cases.

For the values of Q nearest zero, the budget constraint involves paying only the lesser of the two fixed costs. If this option also has the lowest marginal price, as Rx apparently did for many consumers early in the opioid epidemic (especially for those covered by insurance plans with generous copays), then the larger fixed cost would never be paid regardless of Q and the budget constraint would be a single segment, such as the line through allocation B shown in Figure 1a. Otherwise, at greater quantities (e.g., consumers purchasing larger volumes because they have accumulated an opioid tolerance) the budget constraint involves paying the greater of the two fixed costs instead of, or in addition to, the lesser of the two. Either way, the budget set is not convex because it has a boundary with a less steep slope at higher quantities than near $Q = 0$. Overall, the budget constraint could consist of three segments, as shown in the Appendix, or two segments as in Figure 1b.

Which of the three segments contains the utility-maximizing allocation for a consumer with preferences $u(Q,z;1)$ is indicated by comparing the three values $v(p_R,y-f_R)$, $v(p_I,y-f_I)$, and $v(C(1,p_R,p_I;0,0),y-f_R-f_I)$, where v is the aforementioned indirect utility function for a hypothetical consumer facing a linear budget constraint.[10] Figure 1a shows cases with $v(p_R,y-f_R)$ greater than the other two, so that all opioid consumption is Rx consumption regardless of whether Rx prices

[10] Two or three points simultaneously attain the optimum when some of the values coincide.

are high as they are at allocation A or lower as at allocation B. Opioid consumption must fall with Rx prices or be a Giffen good. More surprising is Figure 1b where $f_R < f_I$ and $v(p_R,y-f_R)$ equals either $v(p_I,y-f_I)$ or $v(C(1,p_R,p_I;0,0),y-f_R-f_I)$. Such consumers are indifferent between consuming Rx only (allocation B) and at least some Im (allocation C, where the marginal cost of Q is either p_I or $C(1,p_R,p_I;0,0)$), perhaps mixed with Rx. If consuming at B, their response to a *small increase* $dp_R > 0$ in the Rx price is to consume instead at allocation C, which has *discretely more* total opioids and *discretely less* Rx and all other goods.[11]

This "jump" result for consumers on the margin between two budget segments derives from the convexity of the budget set rather than any assumptions about relative income and substitution effects.[12] A marginal increase in their Rx price induces a discrete substitution effect in exactly the Hicksian sense because the consumer on this margin stays on the same indifference curve. The amount of substitution in the price dimension is either $p_R - p_I > 0$ or $p_R - C(1,p_R,p_I;0,0) > p_R - p_I$ depending on whether the switch is to Im or mixed consumption. This quantitative result and the Roy's-identity properties of v are essential for the results that follow.

Market-level demand.—Figures 1a and 1b illustrate choice by a single type of consumer, but markets consist of many consumers who are heterogeneous in consumption histories, drug tolerance, their cost of participating in illegal markets, and other dimensions. For a simple derivation of price effects, the market-level results that follow take the special case with $f_R = 0 < f_I$ and without income effects or mixed consumption.[13] The preference parameter θ is assumed to be common across consumers and shift demand multiplicatively. The Appendix shows similar aggregation results with income effects, income heterogeneity, and preference heterogeneity.

Let there be a continuum of consumers who differ only in terms of their Im fixed cost f_I. All consumers face the same marginal prices $\{p_R,p_I\}$ and have the same preferences for Q versus other goods. $F(x) \in [0,1]$ is the fraction of consumers with $f_I \leq x$ and the corresponding density

[11] In order for an increase in the Rx price to induce a shift from B to C, rather than the reverse, C must have less Rx consumption than B, as required by the elasticity restriction on $Q(q_R,q_I)$.

[12] Allusions to the result appear in papers about the Peltzman (1975) effect. Higson and Kenkel (2004) note that teenagers, who face higher average prices for alcoholic beverages, are more prone to binge drinking.

[13] Without income effects, the indirect utility function has $\partial v/\partial y = 1$ for all prices and incomes. Despite having $f_R = 0$, mixed consumption ($q_R q_I > 0$) is not optimal if Rx and Im are close enough substitutes in $Q(q_R,q_I)$ in the sense defined in the Appendix.

function is $F'(x) \geq 0$. Let $\theta f^*(p_R,p_I) \equiv \theta[v(p_I,y) - v(p_R,y)]$ denote the critical value of the Im fixed cost that leaves the consumer indifferent between sourcing from Rx and Im. Because income effects on opioid demand are ruled out, both $f^*(p_R,p_I)$ and the price derivative of v are independent of y. The magnitude of the price derivative of v is $H(p) > 0$ with $H'(p) < 0$, so that the consumer's Hicksian demand function is $\theta H(p)$. It follows that $\theta f^*(p_R,p_I)$ is the area under that demand function between the prices p_I and p_R. The fraction of consumers that source from Im rather than Rx is therefore $F(\theta f^*(p_R,p_I))$, with each of them consuming $Q = \theta H(p_I)$. The remaining consumers choose $Q = \theta H(p_R)$.

LEMMA 2 (Market-level comparative statics). Let $D(p_R,p_I,\theta)$ denote aggregate opioid consumption as a function of the two marginal prices and the common demand parameter θ

$$D(p_R,p_I,\theta) \equiv F\big(\theta f^*(p_R,p_I)\big)\theta H(p_I) + \big[1 - F\big(\theta f^*(p_R,p_I)\big)\big]\theta H(p_R) \tag{2}$$

In the neighborhood of $\theta = 1$, the comparative statics of aggregate opioid demand are

$$\begin{aligned} dD(p_R,p_I,\theta)|_{\theta=1} &= \big[1 - F\big(f^*(p_R,p_I)\big)\big]H'(p_R)dp_R + F\big(f^*(p_R,p_I)\big)H'(p_I)dp_I + D(p_R,p_I,1)d\theta \\ &+ [H(p_I) - H(p_R)]F'\big(f^*(p_R,p_I)\big)[H(p_R)dp_R - H(p_I)dp_I + f^*(p_R,p_I)d\theta] \end{aligned} \tag{3}$$

Proof. Totally differentiate (2) and evaluate at $\theta = 1$ to arrive at equation (3). QED

The first line of (3) shows the familiar continuous source-specific substitution effects, which are movements along the demand curve $H()$ at the Im price and at the Rx price, weighted by the fraction of consumers sourcing from each. It also shows the direct effect on market demand of proportional changes in the component demands $\theta H(p_R)$ and $\theta H(p_I)$. The final line of (3) shows effects on total demand of the Rx-Im switching induced by price and demand changes.

Lemma 2's total derivative (3) is the origin of several testable quantitative insights about opioid demand that are derived in this paper by setting to zero one or two of the elements of $\{dp_I, dp_R, d\theta\}$. It also can generate predictions for "long-run" price effects by treating the change $d\theta$ as a response to price changes because current tastes are influenced by past consumption, which itself is a function of past prices. A concrete derivation of such predictions follows in a part of this paper that embeds equation (2) in an overlapping generations model of gateway effects. For now, I refer to comparative statics that set $d\theta = 0$ as "short-run" price effects, as distinct from the gateway extension that also provides "long-run" price effects.

Many opioid policies can be modeled as a change in just one of the two prices. Technological progress in illicit-opioid manufacturing or the treatment of adverse health conditions resulting from intravenous drug use can be modeled as $dp_I < dp_R = d\theta = 0$. Common price reductions $dp_I = dp_R < d\theta = 0$ serve as models of technological or policy changes that reduce the health costs of drug addiction or the financial costs of poor health.

Linking consumption with its composition.—Neither common price changes nor changes in the demand parameter are neutral with respect to the composition of opioid consumption. From equation (3), the switching term for a common price change is $-[H(p_I)-H(p_R)]^2 F'(f^*(p_R,p_I))dp_I$, which is quadratic in the gap $H(p_I)-H(p_R)$ because the gap reflects both the consumption change of an individual that switches and the change in the incentive to switch. The switching term for a demand shift, $[H(p_I)-H(p_R)][v(p_I)-v(p_R)]F'(f^*(p_R,p_I))d\theta$, has a magnitude with almost the same determinants. Both switching terms are zero when the two prices are equal ($p_R = p_I$) or no consumers are on the source margin ($F' = 0$) but otherwise reinforce the continuous terms. A common price reduction or a demand increase must therefore reduce demand at least as much as they would without switching, especially when the two prices are significantly different. Proposition 1 links overall consumption with its composition:

PROPOSITION 1 (Overall consumption and its Rx share change in opposite directions). Assume that $p_R > p_I$ and $F' > 0$. The short-run comparative statics for opioid consumption $D(p_R,p_I,\theta)$ and the Rx quantity share $r \equiv \frac{1-F(\theta f^*(p_R,p_I))}{D(p_R,p_I,\theta)}\theta H(p_R)$ have the opposite sign if (a) $H'(p_I)/H(p_I) - H'(p_R)/H(p_R)$ is sufficiently close to zero and (b) the impulse is any one of the following: $dp_I = dp_R \neq 0 = d\theta$ (common price change), $dp_I \neq dp_R = 0 = d\theta$ (Im price change), or $dp_I = dp_R = 0 \neq d\theta$ (preference change).

Proof. From Lemma 2 and $p_R > p_I$, a common price reduction, an Im price reduction, or an increase in the demand parameter must each increase $D(p_R,p_I,\theta)$. Totally differentiating the definition of r (see the on-line appendix) shows that the Im price decrease and θ increase also reduce r. The semi-elasticity restriction (a) yields the same for a common price increase. QED

With $p_R > p_I$, greater opioid consumption typically increases the potential benefit from paying the fixed cost of sourcing from Im, and vice versa.[14] The only exception is when the greater consumption is resulting from cheaper Rx prices. More precisely, $\theta f^*(p_R,p_I) > 0$ is the potential benefit from switching from Rx to Im that would be offset against the fixed cost f_I. As the area under the demand function $\theta H()$ between the prices p_I and p_R, $\theta f^*(p_R,p_I)$ is increased by either increasing θ, reducing p_I by itself, or by reducing both prices by the same amount. Therefore, the three impulses cited in the proof each induce switching from Rx to Im. The semi-elasticity restriction (a) guarantees that the comparative statics for r are dominated by the direction of the switching. These results would be quite different if $p_R < p_I$.

Sufficient statistics sign Rx price effects.—The short-run behavioral effects of shocks specific to the Rx segment are represented by the special case of equation (3) with $dp_I = d\theta = 0$. Any change $dp_R \neq 0$ that occurs from an initial position of $p_R > p_I$ involves both the usual movement along the demand curve $H'(p_R)dp_R$ as well as a switching term in the other direction as consumers on the source margin switch from Rx to Im that is cheaper at the margin. Proposition 2 identifies sufficient statistics for assessing whether and how much the switching term $[H(p_I)-H(p_R)]H(p_R)F'(f^*(p_R,p_I))dp_R$ dominates, in which case short-run opioid demand increases with prescription prices. This representation also proves useful for deriving predictions for the consequences of other types of price changes, including long-run consequences.

PROPOSITION 2 (Sufficient statistics sign the short-run Rx-price effect). Defining the cross-price elasticity of Im demand $CROSS \equiv \frac{\partial \ln F(\theta f^*(p_R,p_I))}{\partial \ln p_R} > 0$, the arc elasticity $ARC \equiv \frac{1-\frac{H(p_R)}{H(p_I)}}{1-\frac{p_R}{p_I}} < 0$, and the point elasticity $POINT \equiv \frac{H'(p_R)}{H(p_R)} p_R < 0$, the effect of the Rx price on opioid consumption can be signed as:

$$Sign\left[\frac{\partial \ln D(p_R,p_I,\theta)}{\partial \ln p_R}\right] = Sign\left[\frac{1-r}{r} CROSS \frac{ARC}{POINT}\left(\frac{p_R}{p_I}-1\right)-1\right] \tag{4}$$

[14] Opioid consumption and several other consumer behaviors involve a kind of increasing returns, namely that marginal cost falls with the amount consumed (Mulligan 2022).

Proof. Begin with equation (3), taking $dp_R \neq dp_I = 0 = d\theta$. Eliminate $F'(f^*(p_R,p_I))$, $F'(f^*(p_R,p_I))$, $H(p_R)$, and $H'(p_R)$ using equation (3)'s normalization $\theta = 1$ and the definitions of r, CROSS, ARC, and POINT. Factoring out the positive factor ($-r$ POINT) yields (4). QED

The magnitudes of lemma 2's switching terms magnitude depend on the density $F'(f^*(p_R,p_I))$ of consumers on the margin between the two sources and the horizontal distance $H(p_I) - H(p_R)$ between the two allocations in Figure 1b. The density effect is entirely summarized in (4) by *CROSS* while the horizontal distance is summarized by the product of *ARC* and the relative price term. As p_R exceeds p_I by enough, either the switching term dominates or there are no longer any consumers on the margin between the two sources. In other words, p_R reaches a level at which total demand slopes the "wrong" way (more consumption at high prices) even though consumer preferences u satisfy the usual quasiconcave assumptions. The formula (3) with $dp_I = d\theta = 0$ is analogous to the formula for a tax revenue Laffer curve, which also must slope the "wrong" way for tax rates that are extreme enough.

Equation (4) indicates whether the price-induced demand jump is large enough for the switching consumers (the price-gap term) and whether switching consumers are prevalent enough (the r and CROSS terms) to offset the fact that consumers staying with Rx do so with less demand. Aggregate consumption must slope down with p_R in the neighborhood of $p_R = p_I$ because the switching term is zero when the two prices are equal. To the extent that the statistics featured in equation (4) vary over time, across regions, between demographic groups, or between market segments, the magnitude and the sign of p_R's short-run effect varies, albeit predictably. Empirical estimates from one context do not by themselves necessarily indicate even the direction of the effect in other contexts, but can when viewed through the lens of equation (4).

Taking p_I as given, the consumption-minimizing prescription price is above p_I but finite.[15] Because the minimizing p_R sets both sides of equation (4) to zero, it increases with p_I and r. In other words, if p_I and r were to fall with additional illicit supply, Rx policy seeking to minimize consumption must reduce p_R in a greater proportion than p_I did unless the behavioral elasticity term *CROSS ARC*/*POINT* happened to change significantly.

[15] Any Rx-price reduction at or below p_I increases demand because there is no switching in that range. A Rx-price high enough to switch all consumers to Im results in aggregate demand of $\theta H(p_I)$, which exceeds $D(p_I,p_I,\theta)$.

Recall from proposition 1 that outward demand shifts $d\theta > 0$ affect the composition of opioid consumption, which is represented as the Im-intensity term $(1-r)/r$ in equation (4). Outward shifts can occur over the life cycle as addictions or tolerance builds, thereby increasing Im-intensity. Im intensity may also increase with age to the extent that the fixed costs of illicit consumption fall as consumers are less supervised by parents and other adults as they age. Either way, proposition 2 suggests that $\partial D(p_R,p_I,\theta)/\partial p_R$ might be negative in populations of children and youth even while positive in the adult population. Related is the "gateway hypothesis" asserting that opioid addiction begins with Rx consumption during followed by a transition to Im consumption to the extent that opioid addiction took hold at the younger ages.

The gateway hypothesis and long-run price effects can be investigated more rigorously by extending the model (1) and (2) to explicitly distinguish childhood/youth from adulthood, as in proposition 3. In that overlapping-generations framework young individuals consume $Q = H(p_R)$. When consumers become old, they make choices according to the program (1), with their preference parameter $\theta > 1$ determined by their consumption when young. Let $OG(p_R,p_I,\theta)$ denote overall opioid consumption in an economy with equal numbers of young and old:

$$OG(p_R, p_I, \theta) = \frac{F\big(\theta f^*(p_R, p_I)\big)}{2}\theta H(p_I) + \left[\frac{1 - F\big(\theta f^*(p_R, p_I)\big)}{2}\theta + \frac{1}{2}\right] H(p_R) \tag{5}$$

Proposition 3 provides the sufficient statistics for signing the short- and long-run price effects.

PROPOSITION 3 (Sufficient statistics sign the long-run Rx-price effect). Let $\frac{d\ln\theta}{d\ln p_R} < 0$ denote the elasticity of the demand parameter θ with respect to historical Rx prices (i.e., those that prevailed when the current old were young and developing their habits). Defining the Rx share of OG $r_{OG} \equiv \frac{\frac{1-F(\theta f^*(p_R,p_I))}{2}\theta+\frac{1}{2}}{D(p_R,p_I,\theta)} H(p_R)$, the long-run cross-price elasticity $CROSSLR \equiv CROSS + \left[1 + \frac{\theta f^*(p_R,p_I)F'(\theta f^*(p_R,p_I))}{F(\theta f^*(p_R,p_I))}\right]\frac{d\ln\theta}{d\ln p_R}$, the long-run point elasticity $POINTLR \equiv POINT + \frac{1-F(\theta f^*(p_R,p_I))}{1-F(\theta f^*(p_R,p_I))+1/\theta}\frac{d\ln\theta}{d\ln p_R}$, and the long-run Rx-price elasticity as $\frac{d\ln OG(p_R,p_I,\theta)}{d\ln p_R} \equiv \frac{\partial\ln OG(p_R,p_I,\theta)}{\partial\ln p_R} + \frac{\partial\ln OG(p_R,p_I,\theta)}{\partial\theta}\theta\frac{d\ln\theta}{d\ln p_R}$, we have

(a) The short-run Rx-price effect $\frac{\partial \ln OG(p_R, p_I, \theta)}{\partial \ln p_R}$ is signed exactly with the condition (4).

(b) The long-run effect is:

$$\frac{d \ln OG(p_R, p_I, \theta)}{d \ln p_R} = \left[\frac{1-r}{r} CROSSLR \frac{ARC}{POINTLR}\left(\frac{p_R}{p_I} - 1\right) - 1\right](-r\, POINTLR) + (1-r)\frac{H(p_R)}{H(p_I)} \frac{d \ln \theta}{d \ln p_R} \tag{6}$$

(c) The short- and long-run effects can have opposite signs.

Proof. The derivations of each (a) and (b) follow the same steps as proposition 2's proof. An example proving (c) is constructed by selecting sufficient-statistic values that yield a positive but sufficient small short-run effect. QED

Note that each CROSSLR and POINTLR are the sum of their short-run counterparts (proposition 2) and a negative term involving the historical Rx price effect on θ. Proposition 3 shows that the long-run effect is the sum of a sufficient-statistics term, which is the long-run analog of (4), and a term with the same sign (negative) as $\frac{d \ln \theta}{d \ln p_R}$. Although the proposition assumes a strong form of the gateway hypothesis – all opioid habits are formed during the first half of life when Rx consumption is the only option – that is sufficient to show that the short- and long-run consumption effects of Rx prices may have opposite signs.

Im price effects.—Observing the aggregate consequences of any one of the three types of price changes – $dp_R \neq dp_I = 0$, $dp_I \neq dp_R = 0$, and $dp_R = dp_I \neq 0$ – provides a lot of quantitative information about the consequences of the other two. Roy's Identity provides quantitative links between the effects of public policies and technological change that would otherwise appear quite different. Take an increase in Rx regulation (equation (3) with $dp_R > dp_I = d\theta = 0$) as compared to the short-run effect of cheaper fentanyl ($dp_I < dp_R = d\theta = 0$). The switching term from the Im-price change has the same magnitude as the switching term from a Rx-price change multiplied by the ratio of Hicksian demands $H(p_I)/H(p_R)$. If Rx regulation induces a lot of switching, then cheaper Im opioids must have an especially large effect on opioid consumption because a lot of switching reinforces the usual substitution effect.

Regardless of the sign of $p_R - p_I$, the consumption effects of p_I vary predictably with the level of p_R. Propositions 4 and 5 offer results of this type.

PROPOSITION 4 (Price interactions in demand). Evaluated at $\theta = 1$, the cross-price derivative of the aggregate demand function $D(p_R,p_I,1)$ is:

$$\frac{\partial^2 D(p_R,p_I,1)}{\partial p_R \partial p_I} = \left\{ F'\big(f^*(p_R,p_I)\big) \left[\frac{H'(p_R)}{H(p_R)} + \frac{H'(p_I)}{H(p_I)} \right] - F''\big(f^*(p_R,p_I)\big)[H(p_I) - H(p_R)] \right\} H(p_R)H(p_I) \tag{7}$$

Proof. Using equation (3) for an expression for $\frac{\partial D(p_R,p_I,1)}{\partial p_I}$ and then evaluate a partial derivative with respect to p_R. QED

The cross derivative (7) is negative unless the density changes enough (in the right direction) to offset equation (7)'s first term in braces. When used to compare Im price effects across groups with differential access to Rx opioids, proposition 4 predicts that the group facing a greater Rx price will, all else the same, have a more negative Im-price effect than the group facing a lower price. The corollary helps assess the magnitude of this effect in the special case with: no density differences between groups (F''=0), a nonnegative RHS of (4), and a demand curve that is no more elastic at $H(p_R)$ than at $H(p_I)$.

COROLLARY (Bounding price-effect differentials). Let $S_R \equiv \big[1 - F\big(f^*(p_R,p_I)\big)\big]H(p_R)$ denote aggregate Rx opioid consumption. If $F'' \geq 0$, $p_R \frac{H'(p_R)}{H(p_R)} \geq p_I \frac{H'(p_I)}{H(p_I)}$ and $\frac{\partial^2 D(p_R,p_I,1)}{\partial p_R} \geq 0$, then the effect of p_R on the Im-price effect is bounded by (8)

$$\frac{\partial^2 D(p_R,p_I,1)}{\partial p_R \partial \ln p_I} / \frac{\partial S_R}{\partial p_R} \geq -\left(1 + \frac{p_I}{p_R}\right) POINT > 0 \tag{8}$$

where the elasticity POINT is evaluated at $H(p_R)$.

Conveniently, the corollary can be applied without measuring the amount of the Rx-price change, which is challenging when Rx regulations and other factors affect the frictions involved with obtaining Rx opioids for nonmedical use rather than the monetary price itself. The corollary infers the amount of the Rx price change from the change $\partial S_R/\partial p_R$ in Rx consumption.

A stronger and more surprising cross effect is revealed by comparing two groups that have the same preferences $H()$ and θ, the same fixed cost distribution F, and face the same Im price. They differ only in terms of the Rx price, with the "low-cost" group paying p_{LO} and the high-cost group paying $p_{HI} > p_{LO} > 0$. Each group-average opioid consumption is represented by equation (2) evaluated at the prices paid by group members.

PROPOSITION 5 (Group ranks reversed by Im price changes). Fix the Rx prices paid by each group. Assume $\frac{\bar{x}}{\theta} < \lim_{p_I \to 0} v(p_I) - v(p_{LO})$, where $\bar{x} > 0$ denotes the upper support of the common distribution function F.

(a) At any common Im price no less than p_{HI}, the average consumption gap between the groups is $[H(p_{LO})–H(p_{HI})]\theta > 0$, with the low-cost group consuming more.
(b) There exists another common Im price in the interval $(0,p_{LO})$ such that the high-cost group consumes more opioids than the low-cost group.

Proof. (a) At the assumed Im price, neither group has a potential benefit from paying the fixed cost. Therefore, each group's average consumption is on the demand curve $\theta H(p_R)$, which involves more consumption for the low-cost group because $H'(p_R) < 0$. (b) is proved by example, namely any value of p_I satisfying $\frac{\bar{x}}{\theta} + v(p_{LO}) > v(p_I) > \max\left\{\frac{\bar{x}}{\theta} + v(p_{HI}), v(p_{LO})\right\}$ and thereby switching all high-cost consumers from Rx to Im but leaving at least some low-cost consumers sourcing from Rx. Any such value satisfies $p_I > 0$ (the assumed upper-support restriction) and satisfies $p_I < p_{LO}$ by construction. Average consumption is $H(p_I)$ for the high-cost group and in the interval $\left[H(p_{LO}), H(p_I)\right)$ for the low-cost group. QED

In other words, there exists a change in Im prices that reverses the sign of the gap between the two groups' average opioid consumption. Although proposition 5 holds even if the cross derivative (7) is not everywhere negative, propositions 4 and 5 have common intuition. Namely, the lower the price of Rx opioids, the wider the range of Im prices that do not affect an individual's opioid consumption. Moreover, the proof reveals that the rank-reversal result is less likely to occur (that is, requires more of an Im price change) in populations with high fixed costs or low values of the demand parameter, of which the young may be an example.

The final proposition is the strongest result of all, establishing equivalence in direction and magnitude between the effects of a common price change on the composition of opioid consumption and the aggregate consumption effects of changing either one of the prices by itself.

PROPOSITION 6 (Equivalence across price changes). If aggregate Im opioid consumption is denoted as $S_I \equiv F\big(f^*(p_R, p_I)\big)H(p_I)$, then

$$\left.\frac{dS_I}{dp_I}\right|_{dp_I=dp_R,\theta=1} = \frac{\partial D(p_R,p_I,1)}{\partial p_I} \text{ and } \left.\frac{dS_R}{dp_R}\right|_{dp_I=dp_R,\theta=1} = \frac{\partial D(p_R,p_I,1)}{\partial p_R}.$$

Proof. Totally differentiate the definitions of S_I and S_R and compare with (3). QED

A common price change reveals the size and direction of the aggregate effect of both types of *ceteris paribus* price changes, without observing either one of them. Conversely, observing only the aggregate effect of one price change is enough to reveal the effect of a common price change on that segment. If empirical studies find, say, no Rx consumption decline as a result of a policy increasing p_R and p_I by the same amount, then in theory tighter Rx regulations would not reduce overall consumption. By closely linking the effects of seemingly different policies, proposition 6 widens the range of evidence that can inform a specific policy.

Proposition 6 is surprising because increasing two prices results in less switching between opioid sources than increasing just one. However, the magnitude of a switcher's effect on aggregate consumption is less than her effect on Im consumption. Hicksian symmetry applied to the problem (1) guarantees that the two exactly offset (even at the individual level) because $H(p_I)$ is both the Im-consumption change of a switching consumer and the incentive to switch in response to an Im price change, while the gap $H(p_I)–H(p_R)$ is both the incentive to switch in response to a common price change and each switcher's contribution to the aggregate.

III. Quantity, Policy, and Price Measures

Proxies for quantities of opioids consumed, opioid policies, and opioid prices are necessary to test the theoretical predictions. Quantity data is essential, especially because the quantitative statements in some of the propositions are about consumption rather than prices. The price and policy indicators help partition the recent history of opioid markets into distinct phases in terms of the direction of significant price changes.[16]

[16] See also CEA (2019) and Powell and Pacula (2021).

Opioid fatalities.—Annual fatalities by region and demographic group are measured 1999-2021 using the on-line CDC-Wonder tools, sponsored by the Centers for Disease Control and Prevention (CDC), for tabulating every death certificate filed with a U.S. state or District of Columbia (essentially every death in the country). Each death certificate "contains a single underlying cause of death, up to twenty additional multiple causes, and demographic data" (Centers for Disease Control and Prevention, 2022). The tools permit tabulation by any of the thousands of underlying causes, or by selected CDC-defined cause groups such as "Drug-induced causes." Death certificates can additionally be tabulated by any of the thousands of (more specific) multiple causes, such as unintentional heroin poisoning. I select only those records where the underlying cause of death is drug-induced causes, which are primarily International Classification of Disease 10th revision (ICD-10) codes X40-44, X60-64, X85, and Y10-14. I further limit the death records to those where opioids are listed as immediate or contributory causes of death (ICD-10 T codes 40.0/opium, 40.1/heroin, 40.2/other, 40.3/methadone, 40.4/synthetic). I take opium, heroin, and synthetic as illicitly manufactured and the other two T codes as Rx opioids. For the purpose of measuring annual fatality rates by demographic group or for the nation, I use population estimates from the CDC-Wonder tools. The all-race analyses use the nine Census divisions, gender, and three age groups.[17] Solely for estimating shares, any death certificate indicating both Rx and Im opioids is considered both a Rx death and an Im death. Due CDC-Wonder's cell-size limits and that proposition 5 refers only to overall opioid consumption, race comparisons pool Rx and Im fatalities.[18]

I assume that within age and gender and year, the opioid fatality rate is proportional to MGEs consumed so that predictions about consumption are also predictions about mortality within age/gender/year. An earlier version of this paper (Mulligan 2020) showed that, while illicit fentanyl is more MGE intensive, the national changes over time in the MGEs of fentanyl seized by law enforcement (perhaps a consumption proxy) closely follows the number of death certificates indicating fentanyl involved in a drug-induced death.

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[17] The age groups are 0-44, 45-64, and 65+. The regional divisions are shown in Table 1.

[18] For the all-races analyses, which involve the ratio of Rx-opioid deaths to Im-opioid deaths, CDC Wonder's minimum cell-size limits results in only 8 cells (of 1,242) with missing mortality data, and those are limited to the years 2000 and 2001 in the East South Central Region. Race comparisons are conducted for the years 2012-2021, when there is no missing data at the division by gender by age-group by year level after 2013 and only 12 (of 216) missing 2012-13.

The first two columns of Table 1 list Census divisions and their relative populations of whites and blacks. The remaining columns show opioid death rates and illicit shares of opioid deaths separately by subperiod. Death rates increased substantially, although their levels and changes vary considerably across geography. The differential changes are often attributed to differential penetration of illicit opioids by geography. Also note that Census divisions with the greater death-rate increases tend to be those with more blacks relative to whites.

Federal policy database.—Although the OxyContin reformulation receives much attention in the literature, it helps to know its relation with other policies since 2000 that might affect the price or accessibility of opioids for nonmedical use. The model distinguishes Rx policies from Im policies.[19] With the former more numerous, I further partition Rx policies along the chain of production, distinguishing prescribing from consumer effort and expenditure.

As detailed in the on-line appendix, policies were identified from Federal Register final rules and from Department of Justice press releases for the years 2001-19, using the search criteria "opioids." A rule was eliminated if I deemed it insignificant or it set policy unrelated to the price, cost, or availability of opioids, such as a 2011 rule changing the name of an advisory committee. Five rules implemented or significantly changed prior rules, agency documents, or statutes, in which case I located and included those prior policies. The results shown in Table 2 suggest that regulatory and fiscal activity is higher for Rx than Im. In the earlier years, opioid subsidies are created and expanded for patients and prescribers while regulations are relaxed. In about 2010 policies begin to swing in the other direction as the with reformulation (see below) and programs discouraging prescription supply to secondary markets. The results also suggest that enforcement of illicit-drug prohibitions was less of a priority between 2013 and 2016.

Opioid price structure.—The key premises about opioid prices in this paper are that: (i) heroin was significantly more expensive per MGE than Rx opioids in the 1990s, (ii) illicit opioids became cheaper over time, especially since 2013, and ultimately cheaper than Rx opioids, and (iii) beginning in about 2011, Rx opioids became more expensive or difficult to access for nonmedical use due to regulatory and fiscal changes. These hypotheses motivate the

[19] See also Savona, Kleiman and Calderoni (2017) complication of criminology studies of parallel legal and illicit drug markets.

analysis and interpretation of the quantity data, without relying on more precise characterization of prices in opioid markets where participants have strong incentives to avoid being measured.

On premises (i) and (ii), market participants have described a per-dose price gap between heroin and prescription opioids that changed from significantly positive in the 1990s to significantly negative in the late 2010s. Rx opioids were once known as "hillbilly heroin" or "poor man's heroin."[20] Heroin is later recognized to be the cheaper alternative.[21] In a recent survey of opioid addiction treatment patients, "almost all—94 percent—said they chose to use heroin because prescription opioids were 'far more expensive and harder to obtain'."[22]

An earlier version of this paper (Mulligan 2020) finds the year 2013 to be a turning point for both survey reports of ease of heroin access and the share of illegal contraband arriving in crime labs that was fentanyl or heroin. Before then, illicitly-manufactured fentanyl was largely absent from the drug supply with the exception of brief localized episodes that ended with a shutdown of the source by law enforcement. After, consumers frequently received heroin mixed – some would say adulterated – with fentanyl. Fentanyl "is phenomenally inexpensive per dose in wholesale markets" and enough cheaper to largely displace heroin from illicit markets, as it has done in some countries and regions of North America.[23] Likely explanations include technological advances among illicit manufacturers and new smuggling opportunities.

Prescription opioid pills taken for nonmedical use are many times crushed or dissolved so that they could be injected or snorted (contrary to the prescribed method). With this in mind, the Food and Drug Administration (FDA) in 2010 approved new "abuse-deterrent formulation" opioids that could not be abused as easily, thereby increasing the full price of Rx opioids from the perspective of Rx abusers. As shown in Table 2, in 2010 and subsequent years fiscal and regulatory policies would move in the direction of discouraging nonmedical use of Rx opioids.

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[20] Butterfield (2001), Jayawant and Balkrishnan (2005), and Quinones (2015).
[21] Cicero, Ellis and Surratt, et al. (2014), Cicero, Ellis and Kasper (2017), and National Academies (2017).
[22] National Institute on Drug Abuse (2018).
[23] See Pardo (2019, pp. 20, 109, 119, 125) on British Columbia, Estonia, and Latvia. Mortality and NFLIS data suggest that this had also occurred in most of the northeast U.S. by 2019.

IV. Empirical findings by age, geography, and race

Much previous research makes a strong case that substitution between medical and illicit opioid markets has been substantial enough that at times the demand for opioids has been increasing in the price of prescription opioids. This paper shows what the previous findings may reveal about additional consequences of technological and regulatory changes in opioid markets, as presented by propositions 1-6. This section tests those four of the predictions that can be evaluated without precise measures of the magnitude of price changes or data that spans multiple generations. Specifically, the death-certificate data is used to assess whether the sign of the relationship between overall death rate and its composition changed over time; whether opioid deaths fell coincident with the OxyContin reformulation for children and youth, whose opioid consumption appeared to be especially Rx intensive; and whether the opioid death rate for blacks surpassed the rate for whites, especially among middle-aged and older people, during the more recent period when it appears that Im prices fell sharply (premise ii).

IV.A. Segment shares and opioid deaths across areas and over time

If geographic areas differ from each other primarily in terms of illicit-opioid prices, the level of opioid demand at a given price, or both, then the cross-geographic relationship between the overall opioid death rate and its composition depends on the sign of the price difference $p_R - p_I$ present in a typical area (proposition 1). Specifically, to the extent that the admittedly sparse price data suggest that $p_R - p_I < 0$ in many of the years prior to 2010 and $p_R - p_I > 0$ in many of the years after, then the relatively high-consumption areas would be Rx intensive in the early years but Im intensive later. To investigate this, I estimate the linear regression (9):

$$r_{a,d,g,t} = \alpha_{a,g,t} + \beta_t \ln m_{a,d,g,t} + \varepsilon_{a,d,g,t} \tag{9}$$

where m denotes the opioid mortality rate and a, d, g, and t denote age group (0-44, 45-64, 65+), Census division, gender, and year, respectively. Following equation (4), equation (9) denotes the share of opioid deaths involving prescription opioids as r, with the deaths involving both Rx and Im excluded from both numerator and denominator. $\alpha_{a,g,t}$ is a full vector of interaction terms and $\varepsilon_{a,d,g,t}$ is a regression residual. The mortality rate appears as a proxy for opioid consumption, but the log specification together with the interaction vector allows for the fatalities per unit of opioids consumed to vary by age, gender, and year.

Each year t has 53 or 54 age/gender/Census-division cells used to estimate the regression coefficient β_t.[24] Figure 2 shows point estimates and confidence intervals. The cross-area relationship between opioid deaths and their Rx intensity changed from positive to negative in about 2012 or 2013. A sign change sometime during the period 1999-2021 is consistent with the proposition 1, although without more precise price estimates we do not have a prediction as to the exact date. These findings do not rule out alternative explanations for the sign change.[25]

IV.B. The differential effects by age of OxyContin's reformulation

Within months of the FDA's approval of reformulated Rx opioids, prescriptions of OxyContin (the primary Rx opioid that was reformulated at that time) fell sharply (Schnell 2018). Several previous studies of effects of reformulation can now be interpreted and placed into a broader context by using the sufficient statistics results.[26] Viewing their results through the lens of the demand model, equation (4) evaluates to about zero for the overall population in about 2010, but would tend to evaluate to less than zero for groups that are especially Rx-intensive in their opioid consumption, such as children and youth.[27]

The literature on the consequences of OxyContin's reformulation does not estimate separate empirical models for children/youth and adults. The exception is Alpert, Powell, and Pacula (2018), which finds different effects of the reformulation through the year 2013 on the heroin fatality rate among persons aged 0-24. Although their paper does not report whether the reformulation resulted on fewer overall opioid deaths for those aged 0-24, the estimates of overall opioid fatality rates (Rx and Im) shown in my Figure 3 suggest that it may have. The

[24] Each of the 54 cells is weighted by its total number of opioid deaths 1999-2021. For each of the years 1999, 2000, and 2002, one cell (with weight less than 1/800th of the combined remaining 53 cells) is omitted from the regression because the cell had zero opioid deaths. The on-line appendix shows that including the opioid fatality rate in levels rather than logs does not change the basic pattern shown in Figure 2.

[25] By construction, proposition 1 allows for multiple directions of causality. High consumption may cause composition changes in the sense that the underlying impulse is $d\theta \neq 0$ or composition changes ($dp_I \neq 0$) systematically affect the level of consumption.

[26] Ruhm (2019b) concludes that "the release of an abuse-deterrent formulation of OxyContin in 2010 reduced [Rx] demand but almost certainly fueled some substitution to heroin". Mallatt (2018) finds a connection between reformulation and increased heroin crimes in counties that had been OxyContin intensive. Alpert, Powell and Pacula (2018) and Evans, Lieber and Power (2019) find that reformulation reduced Rx deaths and increased Im deaths, leaving total deaths about constant.

[27] Mulligan (2020) finds that the Rx share of opioid deaths falls with age at least until age 18.

rates for minors and adults follow similar time patterns through about 2010. Thereafter economically and statistically significant gaps emerge between the two series.[28] The fact that the minors' series is the only one to remain several years below its 2010 values suggests that, against falling Im prices and other factors increasing fatalities, more strict prescription policies may have reduced fatalities among minors. More work is needed to determine whether reduced opioid use among minors was enough that the long-run effect is negative even if the short-run effect is not.

IV.C. The Rank Reversal of Blacks and Whites

For years, scholars noted that the opioid death rate was significantly less among blacks than whites. Referring to the years since 2000, Case and Deaton (2020, p. 65) report that "blacks were not suffering the epidemic of overdoses, suicide, and alcoholism." Although expecting fentanyl to narrow the black-white gap, Case and Deaton conclude that blacks' prior experience with the 1980s crack epidemic had left a younger generation "disgusted" by the harms of drug addiction. However, to the extent that much of the early 2000s race gap in opioid deaths was due to differential prescription access, the possibility of substitution between Rx and Im has strong predictions for the evolution of the gap over time as illicit opioids became cheaper.[29] Especially, proposition 5 predicts that falling illicit-opioid prices would eventually push the black death rate past the white rate even if blacks and whites paid the same illicit-opioid price, had the same preferences, and had the same within-race distribution F of fixed costs. Less of a price decline is required to reverse the races at older ages when opioid habits and tolerance would be greater.

Figure 4 displays time series of the black-white gap in opioid fatality rates 2012-21. In each cross section of persons alive at the beginning of the year, the gaps are adjusted for gender, age group (0-44, 45-64, and 65+), and Census division by regressing an opioid fatality indicator on indicators for those three variables as well as race, whose rescaled coefficient is the gap shown in the figure. The Census-division adjustment is meaningful because black population shares and the recent increase in fentanyl deaths are positively correlated across areas. When the

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[28] Figure 3's 95 percent confidence interval around the point estimate m_t is $\pm 1.96\sqrt{(1-m_t)m_t/n_t}$, where n_t is the number aged 0-17 in year t. Confidence intervals are not shown for the adult series because they are small on the scale of Figure 3's vertical axis.

[29] A number of studies found lower opioid prescribing rates for black patients as well less health-insurance coverage, both of which may have affected the price and availability of prescription opioids (Rambachan, et al. 2021, Todd, et al. 2000, Lowe, et al. 2001, Pletcher, et al. 2008, Buchmueller and Levy 2020).

sample is limited to persons aged 45+, the adjusted race gap reverses by 2016, whereas as recently as 2021 the black rate remains lower among persons under age 45.[30]

The magnitude of the race gap change predicted by the corollary is significant on the scale of the actual change, although not necessarily explaining all of it. For example, consider a point elasticity of –0.5, illicit prices that fell by a factor of 3 after 2013 after having been equal to Rx prices, and that differential access to prescriptions was enough to put the initial prescription mortality rate gap 4 per 100,000. The inequality (8) predicts a change in the overall opioid death rate for blacks that exceeds the white change by at least 4.4 (per 100,000).[31] By comparison, Figure 4 shows a differential change of 3.9 from 2012 through 2019 and 9.9 through 2021. An alternative explanation for the race reversal is that black rates of non-opioid drug use may have been greater and their non-opioid drugs of choice are increasingly adulterated with fentanyl.[32] The many disruptions of employment, medical care, etc., associated with the COVID-19 pandemic are also relevant. While more research is needed to have a complete quantitative model of black-white differences, the results suggest that the Rx-Im substitution patterns observed in connection with the reformulation of OxyContin may have important similarities to the forces that pushed blacks' death rate past whites.'

V. Extensions

The Appendix shows similar aggregation results with income effects, income heterogeneity, and preference heterogeneity that were left out of the aggregate model featured in propositions 1-6. Even if utility maximization is relaxed to be nonsatiation, price changes still have the two basic effects in the aggregate: a jump from one part of the budget set to another among some consumers and what would appear to be an ordinary substitution effect among those

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[30] Standard errors for the race gap are about 0.1 per 100,000 population, as to be expected given that the black population is about 40,000,000 and most years have an opioid death rate exceeding 2 per 100,000 among blacks.

[31] $4.4 = 4*[(1+1)(-0.5)]*\ln(1/3)$, where the term in square brackets is the numerical counterpart to the right hand side of the inequality (8).

[32] Furr-Holden, et al. (2021). However, adulteration may not be the entire explanation because the race gap in elderly opioid deaths follows a similar pattern to the age 45+ series shown in Figure 4 even though there was no race gap among the elderly before 2013 in terms of non-opioid drug deaths.

who do not jump. Another extension of (2) is to have separate accounting for consumption and deaths, with illicit consumption being more dangerous. Mulligan (2020) analyzes these cases. The model (1) is also a foundation for additional analysis of mixed consumption. Proposition 1 also suggests that addiction treatment programs, to the extent they reduce consumption capital, would not only be associating with less opioid consumption but also a different composition.

A richer lifecycle model would include a Rx-Im choice both before and during adulthood. Although a rigorous definition of aggregate-level epidemic dynamics is beyond the scope of this paper, I note that the model (2) predicts that even a constant trend for the log of prices or for the log of the demand parameter would result in a sudden surge in aggregate consumption as consumers switch from Rx to Im. The peak contribution of switching to consumption growth would be at peak density, thereby giving the appearance of an "epidemic" or "diffusion" even though consumers in my model are not interconnected.

VI. Conclusions

This model in this paper predicts the opioid-consumption effects of a range of policy and technological changes including prescription regulation, technological progress in illicit manufacturing, law enforcement, opioid-overdose treatments, and the labor market opportunity costs of drug addiction to the extent that such policies influence the full price of opioids. Not only do diverse shocks fit into a uniform structure, but the model reveals close quantitative relationships among their effects. Proposition 6 is one such result, establishing an equivalence in direction and magnitude between the effects of a price change common to prescription and illicit opioids (such as overdose treatments) on the composition of opioid consumption and the aggregate consumption effect of changing either one of the prices by itself.

In theory, even the direction of short-run policy effects can change over time, which is consistent with previous empirical findings that increased prescription supply increased opioid mortality during an earlier era whereas later prescription-supply reductions also increased opioid mortality. With its sufficient statistics results (propositions 2 and 3), the model clarifies how the sign and magnitude of the price difference between prescription and illicit markets may be responsible for these changes. The model also shows how recent fentanyl deaths among whites would appear at least partially a consequence of prior prescription habits (proposition 3) at the same time their opioid consumption would be surpassed by blacks, who have little prescription history (proposition 5). The paper presents death-certificate data showing that, in fact, (i) the black-white gap in opioid-fatality rates changed sign, (ii) fatality rates among children and youth diverged from adult rates, and (iii) the cross-area relationship between the opioid fatality rate and its composition changed sign.

This is also the first paper to comprehensively catalog the dozens of changes in federal opioid policy, identified in nineteen years of the Federal Register and from the Department of Justice press releases, that potentially influence prices and costs. The overall pattern revealed in Table 2 is that policies subsidized and facilitated opioid prescriptions from the year 2000 until about 2010. Later Rx regulations were tightened while the war on illegal drugs was relaxed.

Much more can be learned about opioid markets. A significant fraction, if not a majority, of opioid misuse is sourced from illicit markets where the accuracy and variety of price and quantity measures are especially deficient. Such data would be a big step forward toward quantifying the price effects of many of the policies recorded in Table 2. Better predictions would also be possible with estimates of short and long run supply elasticities, and how they are different for heroin and fentanyl. Both the economic model and the policy database would be usefully extended to include drug-treatment policies.

VII. Appendix: Additional Consumer-Theory Results

Budget set properties.—The consumer's budget constraint is piecewise linear in the $[Q,z]$ plane, formed as the upper envelope of the three linear budget constraints (Lemma 1). Assuming that $f_I > f_R \geq 0$, four possible configurations are possible depending how p_R/p_I fits into the interval $0 < Q_{10} < 1 < Q_{01} < \infty$, where Q_{xy} denotes the magnitude of the marginal rate of substitution in $Q()$ evaluated at $q_R = x$ and $q_I = y$. The upper envelope is only one piece if $p_R/p_I < Q_{10}$ (Figure 1a). If income is great enough, the upper envelope consists of two pieces (as in Figure 1b) if $p_R/p_I \geq Q_{01}$ or $Q_{10} < p_R/p_I \leq 1$. If $f_R > 0$, the mixed (Im-only) constraint is dominated by the other two in the former (latter) case, respectively. The remaining interval is where three pieces are possible when $f_R > 0$, with the mixed piece forming the upper envelope at the highest quantities because $C(1,p_R,p_I;0,0) < p_I < p_R$. When $f_R = 0$, the only difference is that the interval $1 < p_R/p_I < Q_{01}$ cannot have three pieces because the Im-only piece is dominated by mixed consumption. The three-piece case is also less likely when Rx and Im are close substitutes: a small gap between Q_{10} and Q_{01}. $Q_{01} \leq \frac{p_R}{p_I} \vee \frac{y-f_I}{y} \leq \frac{C(1,p_R,p_I;0,0)}{p_R}$ are two cases aggregating to equation (2).

Additional heterogeneity.— Let $\theta \in \Theta$ be a vector indexing consumer preference characteristics; the main text is the scalar special case. Average consumption (2) is generalized as $\int_\Theta \int_0^{f^*(p_R,p_I;\theta)} M(p_I, y-f;\theta) g(f,\theta) df\, d\theta + \int_\Theta \int_{f^*(p_R,p_I;\theta)}^{\infty} M(p_R, y;\theta) g(f,\theta) df\, d\theta$, where M denotes the Marshallian demand corresponding to the indirect utility function v, now indexed by θ. $g(f,\theta)$ is the density function. With the following definitions, the effect of p_R on average consumption is still signed by equation (4); see the on-line appendix for details and proof steps.

$$CROSS \equiv p_R \frac{\int_\Theta \frac{A(\theta)}{ARC} \frac{\partial \int_0^{f^*(p_R,p_I;\theta)} M(p_I, y-f;\theta) g(f,\theta) df}{\partial p_R} d\theta}{\int_\Theta \int_0^{f^*(p_R,p_I;\theta)} M(p_I, y-f;\theta) g(f,\theta) df\, d\theta}, ARC \equiv \frac{\int_\Theta A(\theta) g(f^*(p_R,p_I;\theta),\theta) d\theta}{\int_\Theta g(f^*(p_R,p_I;\theta),\theta) d\theta}$$

$A(\theta)$ denotes an individual-level Hicksian arc elasticity and ARC its aggregate among consumers indifferent between sources. POINT is defined as the consumption-weighted average Marshallian point elasticity among Rx consumers. r still denotes the Rx quantity share.

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Consumption responses to Rx price changes

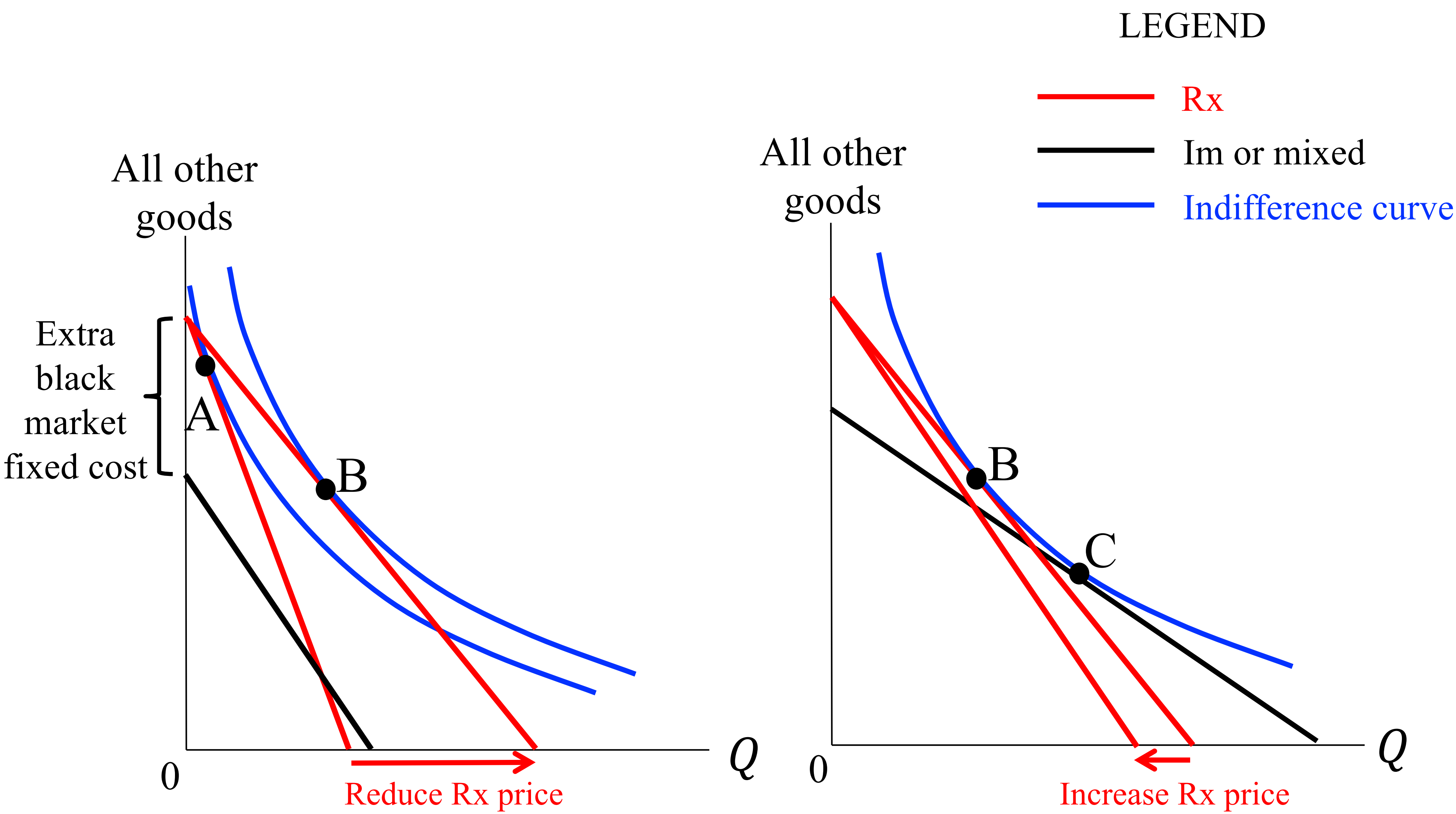
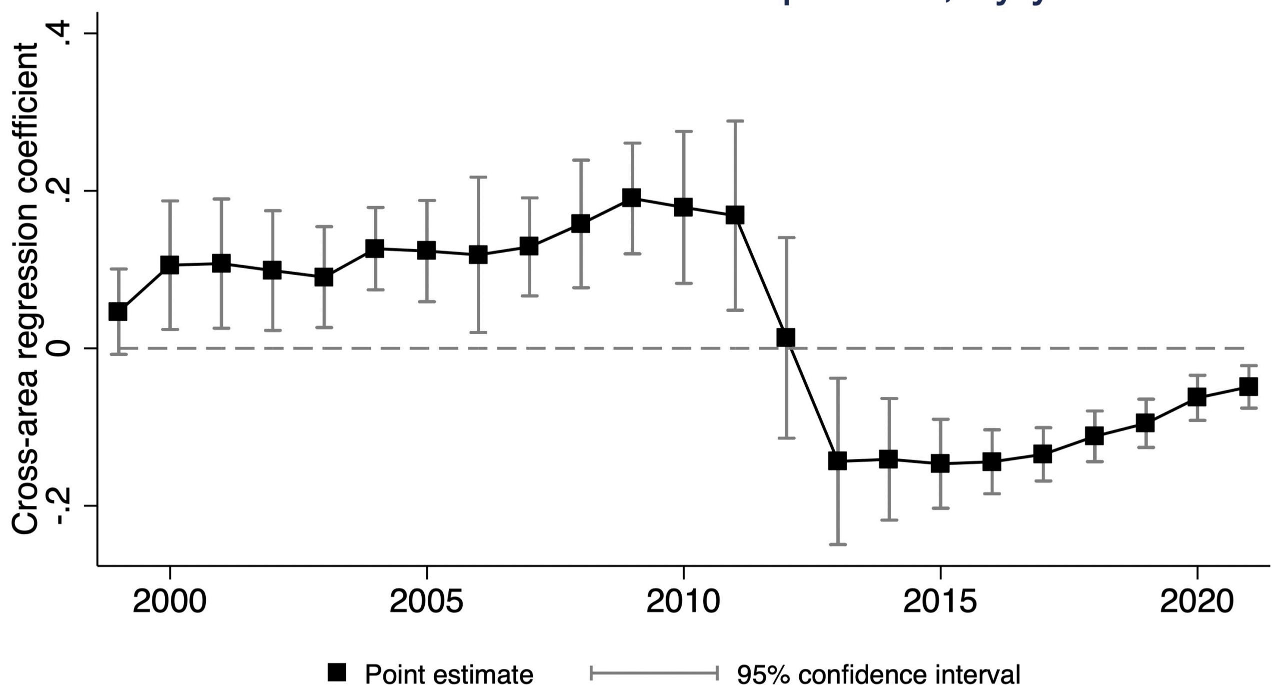


Fig 1a. Im marginal prices are high

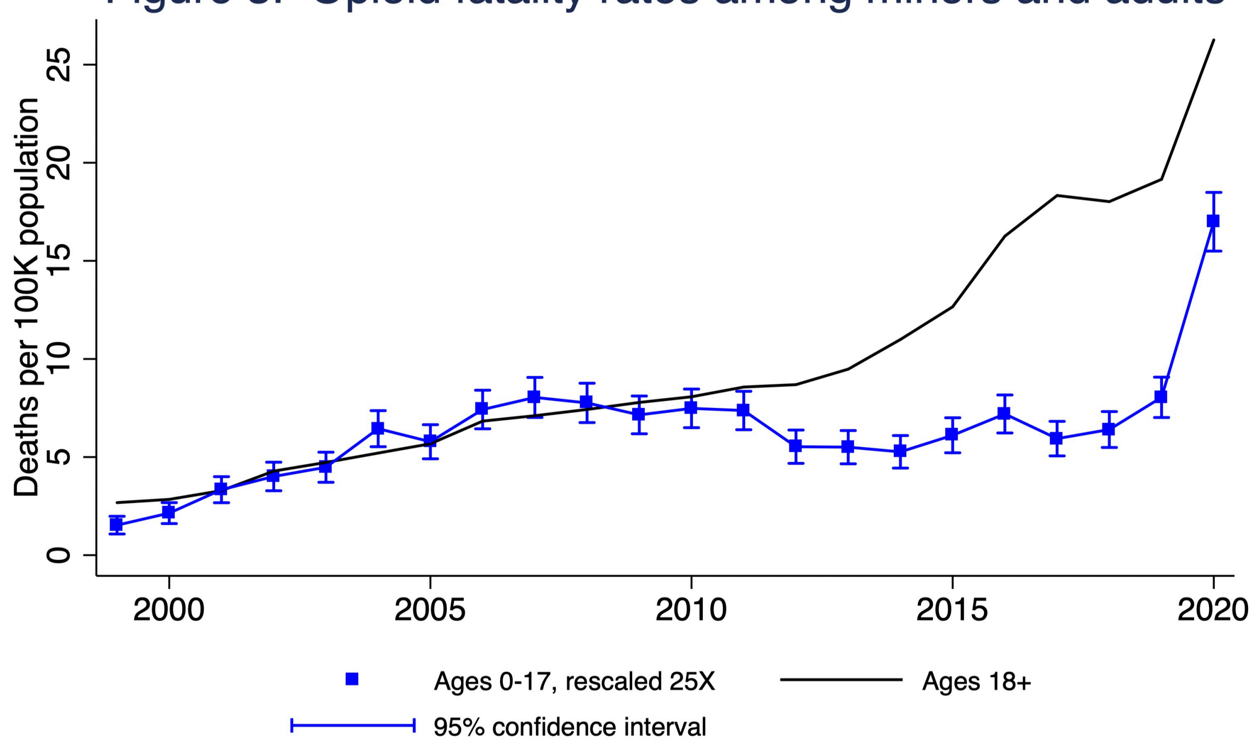
Fig 1b. Im marginal prices are low

Figure 2. The cross-area relation between opioid death rates and their Rx-Im composition, by year

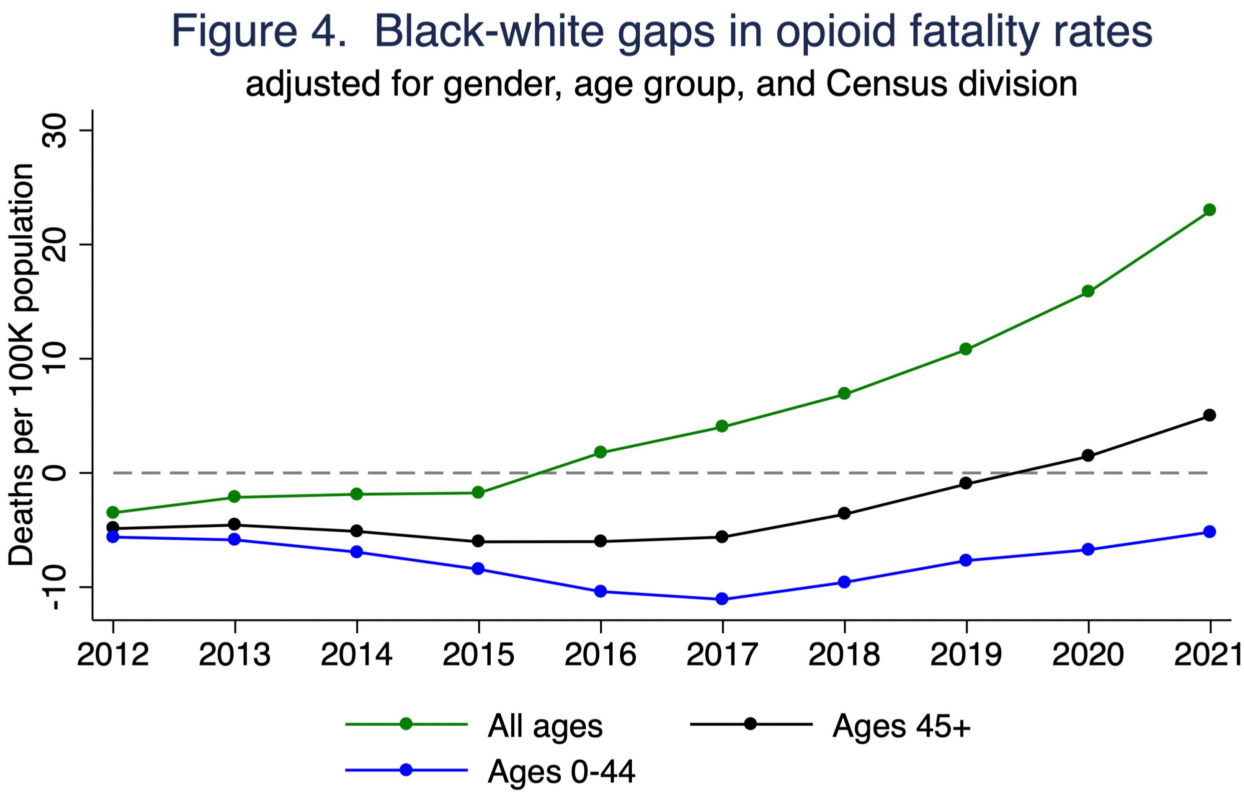


Notes: Observations are age group by gender by Census division, weighted by cumulative opioid deaths 1999-2021.
Each year, the Rx share of opioid deaths by is regressed on gender-age group interactions and log opioid deaths per 100K.
The horizontal axis is shown as a dashed line.
Source: CDC Wonder.

Figure 3. Opioid fatality rates among minors and adults



Note: Death rate for children is multiplied by 25 to show on the same scale with adults.
Adult confidence intervals are not shown because they are less than 0.5 per 100K
Population and mortality source: CDC Wonder.



Notes: For each year an opioid death indicator is regressed on indicators for race, gender, age group, and Census division, with 100K times the black coefficient shown in the figure. A regression observation is a black or white U.S. resident alive January 1. Gap confidence intervals (not shown) are less then 0.5 per 100K population. Population and mortality source: CDC Wonder.

Table 1. Death certificate summary statistics

Census Division	Black pop. %	Opioid deaths per 100K 1999-2012	2013-21	change	Rx % of opioid deaths 1999-2012	2013-21	change, % pts
New England	8%	4.8	24.3	19.5	62%	21%	-40
Middle Atlantic	17%	3.5	17.0	13.6	57%	25%	-32
East North Central	13%	4.0	18.9	14.9	57%	24%	-33
West North Central	8%	3.5	9.2	5.7	65%	33%	-33
South Atlantic	24%	5.6	16.8	11.2	74%	31%	-42
East South Central	21%	4.7	16.7	12.0	81%	35%	-46
West South Central	16%	4.0	6.6	2.5	69%	40%	-29
Mountain	5%	7.1	13.2	6.1	74%	43%	-31
Pacific	8%	4.6	8.4	3.8	70%	37%	-34
U.S.	15%	4.6	14.1	9.5	68%	30%	-38

Notes: Black population is percentage of black+white. Death certificates indicating both Rx and Im opioids count as both for the purposes of calculating percentages. Source: CDC Wonder.

Table 2. Changes in Federal incentives related to the market for opioids

	Incentives for:		
Year	Prescribers	Patient Rx purchases	Illicit Manafucture
2000	VHA mandates "5th Vital Sign"[a]		
2001	Pain management becomes part of Medicare/Medicaid accreditation (CMS delegated to TJC)[b]		
2005	DEA clarifies that opioid refills are not permitted, but that subsequent prescriptions can be obtained without appointment.[c]		
2006		Medicare Part D begins covering opioids, but not benzos (CMS)[d]	Fentanyl manuf. shutdown; DEA prohibitions follow.[e]
2007	DEA allows multiple prescriptions with a single office visit.[f] CMS publicizes & requires quality measures, including HCAHPS pain questions, for full reimb.[g]		
2010	DEA allows electonic Rx.[h]	First DEA Rx take-back programs.[i] Product reformulation and withdrawal (FDA)[j]	
2012	CMS penalizes low HCAHPS scores.[k]		
2013	VHA Opioid Safety Initiative; peak VHA opioid Rx[l]	Medicare Part D begins covering benzos too (CMS).[m]	Holder memo: DOJ does not prosecute nonviolent drug crimes[n]
2014	DEA switches Hydrocodone combination products from Schedule III to Schedule II.[o]	Medicaid expansion; deadline for other insurance to cover benzos. (ACA)[p]	
2017	CMS changes its use of pain management surveys.[q]	FDA first requires benzos to carry an opioid-interaction warning.[r]	Holder memo reversed.[s]
2018	Rx quotas tightened.[t]	SUPPORT Act[u]	SUPPORT Act[u]
2019	CMS removes pain management questions from HCAHPS[v]		Series of new DEA prohibitions.[w]

Notes:

[a]Department of Veteran Affairs (2000)

[b]Joint Commission on Accreditation of Healthcare Organizations Pain Standards for 2001. See also 66 FR 4076.

[c]DEA. Clarification of... Prescribing Schedule II Controlled Substances. August 2005.

[d]70 FR 4228 (January 2005).

[e]DEA prohibits fentanyl ingredients in 2008 (73 FR 43355) and 2010 (75 FR 37295).

[f]DEA. Issuance of Multiple Prescriptions for Schedule II Controlled Substances. Nov 2007.

[g]71 FR 68193 (November 2006).

[h]DEA. 75 FR 61613 (October 2010)

[i]DEA. "DEA Heads First-ever Nationwide Prescription Drug Take-back Day."

[j]https://www.medpagetoday.com/productalert/devicesandvaccines/19409 and https://www.fda.gov/drugs/drug-safety-and-availability/fda-drug-safety-communication-fda-

[k]CMS. Medicare Program; Hospital Inpatient Value-Based Purchasing Program 76 FR 26493

[l]Good (2018).

[m]77 FR 22076 (April 2012).

[n]Holder, Eric. "Department Policy on Charging Mandatory Minimum Sentences...."

[o]DEA. Schedules of Controlled Substances: Rescheduling Hydrocodone Combination Products from Schedule III to Schedule II. August 2014. 79 FR 49661

[p]Benzo coverage is in Section 2502 of the Patient Protection and Affordable Care Act.

[q]Effective Oct 2017, the pain part of HCAHPS would no longer be used for VBP, although still for accreditation (81 FR 79571). Effective Oct 2019, outpatient departments would participate in their version of HCAHPS (OAS CAHPS; 71 FR 79771).

[r]https://www.fda.gov/drugs/information-drug-class/new-safety-measures-announced-opioid-analgesics-prescription-opioid-cough-products-and

[s]Sessions, Jeff. "Department Charging and Sentencing Policy."

[t]83 FR 32784 (July 2012).

[u]SUPPORT criminized possession of controlled-substance analogs, restricted illicit import, and encourage unused Rx disposal.

[v]83 FR 58818, with hospitals to first administer with January 2022 discharges.

[w]Spanning 5/2016 through 11/2019, 11 DEA rules put various fentanyl analogs on Schedule I.