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PRICES AND FEDERAL POLICIES IN OPIOID MARKETS

Casey B. Mulligan

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ABSTRACT

Consumer theory is extended to incorporate deviations from the law of one price that are common in markets for prescription and illicitly-manufactured opioids. The extension helps to resolve "puzzling" findings in the literature, such as race and age gaps in mortality rates and a failure of increases in the full price of Rx opioids to reduce opioid fatalities. The theory also identifies characteristics of public policies that are essential for predicting consumer behavior and thereby population life expectancy. Most surprising is that, with heroin and fentanyl relatively cheap of late, any Rx opioid policy could – and likely does – have the opposite total-consumption effect after 2013 than it would before, especially when the more expensive Rx opioid products are differentially affected. The theoretical framework also guides assembly of a dataset of federal opioid policies and assessment of the role of technological change and law enforcement in illicit markets.

Casey B. Mulligan
University of Chicago
Department of Economics
1126 East 59th Street
Chicago, IL 60637
and NBER
c-mulligan@uchicago.edu

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 drops 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 purpose of this paper is to show how the price gap between medical and illicit markets determines the effects of specific opioid policies as well as important interactions between policies.

Although measuring prices in illicit markets is subject to significant measurement error, it is generally understood that that prescription opioids were once "poor man's heroin" (U.S. Department of Justice, National Drug Intelligence Center, 2001) but now "heroin is cheaper and easier to get than prescription opioids" (National Institute on Drug Abuse, 2018a).[2] In a setting like this where potentially close substitutes coexist at substantially different prices, the traditional economic approach delivers conclusions contrary to conventional wisdom. A naïve application of the law of demand would suggest that opioid consumption would be discouraged by increasing the price of one of the opioid alternatives. However, a price increase in one segment might induce consumers to incur a fixed cost that allows them to access a lower marginal price alternative, which encourages them to consume more. Because so many public policies are specific to one alternative or another, it follows that the consumption and mortality effects of one of them cannot be understood, even approximately, in isolation from the others.

[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.

The economic literature has a good start on measuring effects of particular opioid policies at specific points in time. Alpert, Powell and Pacula (2018) and Evans, Lieber, and Power (2019) look at changes in prescription products intended to reduce opioid misuse. Buchmueller and Carey (2018) and Meinhoffer (2018) measure effects of state prescription-drug monitoring programs. Behavioral effects of naloxone, a drug that can reverse opioid overdose, are estimated by Doleac and Mukherjee (2018) and Rees, et al. (2019). Insurance coverage effects are estimated by Zhou, Florence and Dowell (2016), Soni (2018), Powell, Pacula, and Taylor (2017), and Council of Economic Advisers (April 2019). Advertising and internet access are examined by Jena and Goldman (2011) and Nguyen, Bradford and Simon (2019). See also the opioid studies surveyed by Maclean et al (2020). Savona, Kleiman and Calderoni (2017) assemble criminology studies of parallel legal and illicit drug markets.

The idea that the effects of policies depend on the status of illicit markets appears already in this literature, but this paper contributes by showing how many individual policies and possible effects fit together in a simple overall structure. Doing so also reveals a health-economics application of the Lucas (1976) critique: there may be no such thing as "the effect" of a specific opioid policy because even its direction depends on other policies toward substitutes and complements. Even some of the racial disparities that appear large in the raw data can be understood in terms of a common economic framework.

Applying and extending the standard consumer theory is especially valuable in a market like opioids where data is sparse so that policy analysis rationally puts more weight on potentially relevant lessons from other contexts and industries. Section II of this paper provides such a conceptual framework, concluding that the effects of prescription policies, even directionally, depend on policies toward the supply of illicitly-manufactured opioids. The framework yields sufficient statistics that help predict which policies would increase fatal overdoses, which policies would reduce them, and when. By assembling a database of federal opioid policies organized in this way, Section III adds to the available data and reveals distinct policy phases. Section IV presents price and quantity measures, with careful attention to the significant challenges in measuring illicit activity and their relationship with the policy phases identified in Section III. Additional quantity measures shed some new light on the degree to which opioid fatalities are driven by opioid consumption as opposed to changes in fatalities per unit consumed.

The theoretical framework is consistent with both habit formation – that prescriptions can be a gateway to consumption of heroin or fentanyl – and strong substitution between medical and illicit markets in the long run. The empirical age and race patterns shown in Section V, as well as previous findings on the surprising effects of reformulating the leading prescription-opioid brand, suggest that both effects are empirically important. Section VI provides evidence of dramatic increases in illicit supply after 2013 that appears to have created a situation in which total opioid consumption would respond the "wrong" way to prescription prices. Section VII is a quantitative exercise that leans harder on the illicit-market measures but demonstrates how to generalize estimates of behavioral responses during a previous era when illicit opioids were expensive to the more recent era when they are particularly cheap. The final section concludes.

II. Opioid Policies and the Consumer Budget Set

The model has strictly quasiconcave preferences $u(Q,z)$ over two composite commodities: opioids Q and "all other goods" z. Although suppressed in this notation, the relative preference for these two composites may vary with the amount of opioid consumption in the past, as it does in models of habit, addiction, and drug tolerance. The rate of exchange between the composite commodities is the full price of opioids, which includes not only the out-of-pocket cost but also consumer time, effort, hassle, or stigma. Although the nonlinearity of the budget constraint is essential to what follows, the indirect utility function $v(p_Q,y)$ and Hicksian demand function $H(p_Q)$ – for a hypothetical consumer with preferences u and facing a linear budget constraint $y = z + p_Q Q$ – illuminate the derivations.[3] The focus of this paper is how the composite Q is produced and how its full marginal price p_Q varies with the component prices.

[3] Following Becker's (1962) approach to "irrational" choice, Appendix I uses the nonlinearity of the budget constraint and nonsatiation to derive a special case of the market-level results that follow, without relying on utility functions or indifference curves.

II.A. A Household Production Approach to Opioid Choices

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 so that $Q(1,0) = 1$, 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.[4]

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) of obtaining it, 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. Because they quantify the cost of participating in an opioid market net of the cost of participating in a market for another good, the fixed costs can be either positive, negative, or zero. Particularly relevant for opioid markets is the difference $f_I - f_R$, which I expect is often (but not

[4] The health problems include serious and sometimes fatal infections (Collier, Doshani and Asher 2018, Powell, Alpert and Pacula 2019).

always) 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 is itself a fixed cost that may be necessary to access a low quality-adjusted price. Depending on market conditions, Im opioids may also have a lower marginal price per morphine equivalent than Rx opioids.[5] Not surprisingly, the fixed costs allow for persistent deviations from the law of one price: $q_R = 0$ consumers coexisting in the market with $q_I = 0$ consumers even while p_R deviates significantly from p_I.

Let $E(Q,p_R,p_I;f_R,f_I)$ denote the minimum cost of achieving output Q given the fixed and marginal prices of Rx and Im. This cost function partitions the consumer's decision problem in two stages. In one stage, the consumer decides how to produce Q from Rx and Im, which is the minimization that defines E. In the other stage, the consumer allocates his income y between opioids Q and all other goods according to his preferences $u(Q,z)$ subject to the constraint that $z + E(Q,p_R,p_I;f_R,f_I)$ does not exceed his income. Note that the cost function E depends on the shape of Q but not the shape of u. In the absence of fixed costs, the cost is proportional to Q with proportionality factor $E(1,p_R,p_I;0,0)$.

The fact that the consumer's problem separates into stages is also useful for considering intertemporal issues. The amounts of Q consumed in the past affects the willingness to pay for Q in the present. A forward-looking consumer would also be concerned that his present consumption Q contributes to habits that will shape his choices in the future. To analyze intertemporal choice we just extend the utility function u to reflect a time series for Q with habits, addiction, and drug tolerance reflected in the degree to which Qs at various dates are complements (Becker & Murphy, 1988; Pollak, 1970). The opioid consumption Q at any date t must be financed with expenditure $E(Q,p_R,p_I;f_R,f_I)$, where the fixed and marginal prices are specific to date t and the function E embeds the sourcing decision for date t opioid consumption.

[5] 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 itself lowers the marginal price of opioids because each morphine-gram of opioids is more "effective" when delivered that way. On the other hand, illegal sellers forgo some economies of scale in order to avoid detection by law enforcement.

Part of the consumer's minimization problem embedded in the cost function E 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. In each of these cases, the marginal cost of opioids, $\partial E/\partial Q$, is constant. In other words, the consumer's budget constraint is piecewise linear in the $[Q,z]$ plane, formed as the upper envelope of the three linear budget constraints corresponding to the three possible decisions regarding fixed costs: $y = z + f_R + Q\, p_R$, $y = z + f_I + Q\, p_I$, and $y = z + f_R + f_I + Q\, E(1,p_R,p_I;0,0)$, respectively. For the values of Q nearest to zero, the budget constraint involves paying only the lower of the two fixed costs. If this option also has the lesser 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 a morphine tolerance) the budget constraint involves paying the greater of the two fixed costs instead of, or in addition to, the lesser of the two fixed costs. 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 Appendix I, or two segments as in Figure 1b.

II.B. Why Opioid Consumption Might Increase with Prescription Prices: Sufficient Statistics

In Figure 1a, all opioid consumption is Rx consumption regardless of whether Rx prices are high as they are at allocation A or lower as at allocation B. Barring the Giffen case, opioid consumption must fall with Rx prices. The surprising result from Figure 1b is that Rx consumers near the margin with Im, such as consumers with the indifference curve shown in the figure, respond to a *small increase* $dp_R > 0$ in the Rx price by consuming *discretely more* total opioids and *discretely less* of all other goods. The small price change dp_R "jumps" such consumers' choices from allocation B to allocation C, which has less Rx consumption but more total opioid consumption (much of it Im).[6] This result for consumers on this margin derives from the convexity of the budget set rather than any assumptions about relative income and substitution effects or from Figure 1b's static setup.[7] A marginal increase in their Rx price induces a discrete substitution effect in exactly the Hicksian sense because by definition the consumer on this margin stays on the same indifference curve. The magnitude of the substitution effect in the price dimension is either $p_R - p_I > 0$ or $p_R - E(1,p_R,p_I;0,0) > p_R - p_I$ depending on whether the Rx consumer switches entirely to Im or switches to mixed consumption. As will become clearer in what follows, this quantitative result is essential for understanding recent changes in opioid markets.

Figures 1a and 1b illustrate the choice of a single type of consumer, but of course the market consists of many consumers who are heterogeneous in several dimensions, including but not limited to their consumption histories, drug tolerance, and their cost of participating in illegal markets. In order to show a simple derivation of the price effect on aggregate consumption, what follows is the special case without income effects, mixed consumption, or f_R different from zero. All heterogeneity in this case is in terms of f_I and therefore in terms of the propensity to source from Im. All consumers face the same marginal prices (p_R, p_I) and have the same preferences for Q versus other goods. Let $F(p_R,p_I) \in [0,1]$ denote the fraction of consumers that source from Im rather than Rx. F reflects the indifference condition $v(p_R,y) = v(p_I,y-f_I)$ for the consumers on the margin between Rx-only and Im-only, with the value of F as the fraction of consumers with f_I low enough to be on the Im side of that margin. Roy's identity applied to the indirect utility function v implies that each of F's price derivatives is the product of (i) the corresponding conditional demand level (simply $H(p_R)$ or $-H(p_I)$) and (ii) the density of consumers at the value of f_I that makes them indifferent between the two sources.

[6] 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. Also note that the reverse is ruled out by the restrictions on the elasticity of factor substitution in Q.

[7] Allusions to the result appear in the literature on alcohol consumption. Higson and Kenkel (2004) note that teenagers, who face higher average prices for alcoholic beverages, are more prone to binge drinking. Figure 1b could also be applied to the seatbelt studies inspired by Peltzman (1975), although those studies do not explicitly consider the case of a driver who is indifferent between wearing a seatbelt or not. Bell, Ho, and Tang (1998)'s study of fixed shopping costs emphasizes deviations from the law of one price, but it holds fixed the quantity to be purchased by the consumer.

A somewhat more general setup would have consumers also differing in terms of a demand shifter θ, which may depend on amounts consumed in the past or personal characteristics and experiences that affect demand. The same results apply, except that the share function F reflects the cross-sectional distribution of f_I/θ rather than f_I by itself. The consumers sourcing from Im would be those with high levels of opioid demand, low fixed costs, or some combination thereof.[8] Appendix I shows that the signs and marginal rate of substitution in F are the same even with many dimensions of heterogeneity, except that the demand levels quantifying F's marginal rate of substitution are averages across heterogeneous marginal consumers. These Roy properties of F derive from the indifference condition $v(p_R,y) = v(p_I,y–f_I)$ for the consumers on the margin between the two sources, which is the foundation of Appendix I's model as well as the special case featured here.

With these definitions of H and F, aggregate opioid consumption D and its Rx-price derivative are, respectively:

$$D(p_R, p_I) = F(p_R, p_I)H(p_I) + [1 - F(p_R, p_I)]H(p_R) \tag{1}$$

$$\frac{\partial D(p_R, p_I)}{\partial p_R} = [H(p_I) - H(p_R)]\frac{\partial F(p_R, p_I)}{\partial p_R} + [1 - F(p_R, p_I)]H'(p_R) \tag{2}$$

where assuming existence of F's partial derivatives implicitly rules out mass points in the distribution of fixed costs. If $p_R \leq p_I$ and $f_I - f_R > 0$, then all opioids are optimally sourced from prescriptions and total demand is simply $D(p_R, p_I) = H(p_R) \geq H(p_I)$. With $F(p_R,p_I)$ and its price derivatives at zero in this case, the first term on the RHS of equation (2) is zero because marginal changes in the Rx price do not cause any consumer to source from Im instead.

Aggregate behavior can be quite different when $p_R > p_I$ because the first "jump" term has the opposite sign as the final term. It is a special case of a composition effect in which Roy's identity and other features of demand theory restrict the relationship between share changes and the gap between groups. Specifically, the jump term's magnitude depends on the density of consumers on the margin between the two sources and the horizontal distance $H(p_I) - H(p_R)$

[8] In their study of crack cocaine, Galenianos and Gavazza (2017) find that frequent customers are the ones obtaining the lower quality-adjusted prices.

between the two allocations in Figure 1b, which is greater in the mixed case where *Q*'s low price is $E(1,p_R,p_I;0,0) < p_I$. As p_R exceeds p_I by enough, either the jump 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 even though consumer preferences *u* satisfy the usual quasiconcave assumptions. Indeed, the formula (2) 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. This paper identifies sufficient statistics for assessing whether and how much opioid demand increases with prescription prices and applies them to U.S. opioid markets.

A few consumers may have $f_I - f_R < 0$, and thereby might purchase Im opioids even in years when their marginal price is higher. Regardless, the formula (2) has a sufficient statistic format (3) determining its sign. To derive (3), divide both sides of equation (2) by Rx demand and then eliminate the levels and price derivatives of *F* and *H* using the definitions of shares and price elasticities:

$$Sign\left[\frac{\partial \ln D(p_R,p_I)}{\partial \ln p_R}\right] = Sign\left[\frac{1-r}{r} CROSS \frac{ARC}{POINT}\left(\frac{p_R}{p_I}-1\right)-1\right] \tag{3}$$

where $r \in (0,1)$ is the Rx quantity share evaluated at marginal prices, CROSS > 0 is the cross-price elasticity of aggregate Im demand with respect to the Rx price, ARC < 0 is the arc elasticity of *H* between the two prices, and POINT < 0 is the point elasticity of *H*.[9] Equation (3) also holds in the more general framework with income effects and many dimensions of heterogeneity.[10] Either way, Equation (3) is an indicator of whether the jump is large enough for the marginal consumers (the price-gap term) and whether marginal consumers are prevalent enough (the *r* and CROSS terms) to offset the fact that consumers staying with Rx do so with less demand.

Because the first and second terms on the RHS of equation (3) derive from the first and second terms of equation (2), respectively, it is easy to see results of extending this framework to

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[9] CROSS and POINT are the local elasticities of $F(p_R,p_I)$ and $H(p_R)$, respectively, with respect to p_R. ARC is a ratio of percentage gaps between Rx and Im, using Im as the base. The result of making the elasticity substitution is the RHS of (3) times a positive factor ($-r$ POINT).

[10] In that framework, (i) equation (3)'s sufficient statistic POINT represents the Rx-consumption weighted average of individual Marshallian point price elasticities and (ii) ARC represents the Im-consumption weighted average, among marginal consumers, of individual Hicksian arc elasticities (see Appendix I).

account for deaths rather than consumption under the hypothesis that Im opioids are more dangerous per MGE than Rx opioid consumption. Deaths would put more weight on the first (Im) term in equation (2), and therefore more weight on the first term in equation (3). Although source-specific mortality is implicit in the household choice production framework set forth above, it may also be interesting to incorporate it more explicitly in the choice analysis, as in Mulligan (2020).[11]

Aggregate consumption must slope down with p_R in the neighborhood of $p_R = p_I$. To the extent that the statistics featured in equation (3) vary over time, across regions, between demographic groups, or between market segments, the magnitude and the sign of p_R's mortality effect varies, albeit predictably. As noted in my introduction and examined more closely in Section IV, p_R/p_I has especially changed since the 1990s, so that prescription policies that would have reduced mortality before 2013 may be increasing them after. Therefore, empirical estimates of p_R's mortality effect at a point in time do not by themselves generalize to other points in time where equation (3)'s sufficient (and necessary) statistics are different, although they can when combined with equation (3).

Taking p_I as given, the consumption-minimizing prescription price is above p_I but finite. Because the minimizing price sets both sides of equation (3) to zero, it increases with p_I and r. In other words, as p_I and r have fallen in recent years with additional illicit supply, the consumption-minimizing prescription price also fell in a greater proportion than p_I did unless the behavioral elasticity term CROSS ARC/POINT happened to change significantly.[12] This conclusion has not yet been recognized among policymakers, who can affect prescription opioid prices with regulations and subsidies but sometimes assert that prescription policy should be set without regard to illicit markets (Food and Drug Administration, Center for Drug Evaluation and Research, 2017, p. p. 182).

[11] To the extent that source-specific mortality does not affect the second-derivative properties of consumer choice, it can be reflected in the full prices. Second derivative effects are part of the household production function, although I still assume that "all other goods" z can be treated as a composite commodity without regard for how opioid consumption is composed.

[12] By contrast, to the extent that the Hicksian demand H is price inelastic, Roy's Identity says that the prescription price change required to maintain the indifference of marginal consumers as to their opioid source is proportionally less than the change in p_I. In other words, demand-minimizing prescription pricing likely pulls consumers into the prescription market from the illicit market as the illicit market becomes cheaper.

Even though the short-run demand curves are likely different for individuals with drug addictions as compared to new users, equation (3) suggests that both groups could have opioid demand that slopes the "wrong" way with respect to p_R. Having built up tolerance over time, addicts likely have higher levels of demand and therefore greater pecuniary benefits from switching to a source with lower marginal costs. New users with low demand may be more likely to reduce their Rx consumption, or quit opioids altogether, when prescriptions become more expensive.[13] On the other hand, the price elasticity of an addict's demand is also closer to zero in the short run because of the difficulty of reducing drug consumption (Becker & Murphy, 1988), which means that switching does little to increase total consumption. In terms of equation (3), the more price sensitive behavior that characterizes new users (and perhaps all users in the long run) has offsetting effects because one elasticity (ARC) is in the numerator while the other elasticity (POINT) is in the denominator. The offset is substantial, although the effect of POINT is ultimately greater.[14]

II.C. The Substitution Effects of Illicit Prices

As long as marginal prices are no greater for Im than for Rx, Im prices reduce total opioid consumption, which is qualitatively consistent with a naïve application of the law of demand. However, the magnitude of the effect is more surprising because marginal prices increase with the Im price more than one for one. An econometric analysis conducted without regard for the switching between Rx and Im might mistakenly attribute the extra increase in the quantity consumed to a shift in demand that must be explained by preference and other demand factors.

Specifically, consumers who shift from Im to Rx see their marginal price increase by the gap between the Rx and Im prices. With the jump in his marginal price, the marginal consumer's quantity consumed jumps in proportion to his Hicksian price elasticity of demand

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[13] As in the tobacco demand literature, the source-conditional Hicksian demand curve $H(p)$ can represent both the duration of quit attempts and the likelihood of quitting entirely.

[14] If, for example, POINT has the same value everywhere on the arc, then the ratio ARC/POINT decreases with POINT. At a price ratio $p_R/p_I = 2$ and POINT $\in [-1,0)$, the elasticity of ARC/POINT with respect to point is between -0.347 and -0.307. The elasticity closer to zero with linear demand.

and the log difference in the marginal prices of Rx and Im. As the Im analog to equation (2), equation (4) includes both this jump term as well as the more familiar within-source substitution effect $H'(p_I)$:

$$\frac{\partial D(p_R, p_I)}{\partial p_I} = [H(p_I) - H(p_R)]\frac{\partial F(p_R, p_I)}{\partial p_I} + F(p_R, p_I)H'(p_I) \tag{4}$$

The cross-sectional distribution of the gap in fixed costs, $f_I - f_R$, may differ between demographic groups. Group-specific versions of equations (1)-(4) are derived by integrating the fixed cost distribution within groups rather than for the entire population. Obviously, groups with different average fixed costs would have different propensities F and $1-F$ to source from Im and Rx, respectively, as blacks and whites have differed since about the year 2000 (Alexander, Kiang, & Barbieri, 2018). More surprising is that, according to the model, two groups with different fixed-cost distributions would have predictably different dynamics as Rx and Im supply conditions evolved over time. Suppose for the moment that whites and African Americans were the same in terms of marginal prices, the density of consumers on the margin between the two sources, and their source-specific demand functions $H()$. A reduction in p_I, or an increase in p_R, would increase the black-white gap in total opioid fatalities because the two groups have the same jump terms in equations (2) and (4) while blacks have been less prescription-intensive in their fatalities.[15] These counterintuitive price effects on the racial composition of total opioid fatalities would, in theory, be particularly pronounced in the age/sex groups that begin with larger race gaps in the prescription-intensity of their fatalities.

This result for total fatalities is more ambiguous if, additionally, p_R were greater, and the density of marginal consumers less, for blacks, although these two additional differences tend to offset each other in terms of creating a race gap in the jump terms. However, we have another prediction in this case that derives from the racial difference in the composition of the jump terms in equations (2) and (4). While the two races might have roughly comparable jump terms, for blacks it would be a larger jump for a smaller fraction of their population. Rx prices would

[15] Formally, the effect of p_I on black demand relative to white demand is the subtraction of the white version of equation (4) from the black version: $[F_{black}(p_R, p_I) - F_{white}(p_R, p_I)]H'(p_I)$. Here the jump terms are canceled due to the common-density assumption and the restrictions on F imposed by demand theory. Subscripts indicate race for those functions assumed to vary by race.

increase per capita Im consumption more for whites than blacks, even without necessarily changing the race gap for total opioid consumption because it also reduces Rx consumption more for whites. At the same time, lower Im prices would increase total consumption more for blacks than for whites.

Although these results refer to a dichotomy between Rx and Im, similar results could be obtained by subdividing the Im category. Illicitly-manufactured fentanyl and its analogs ("fentanyls") have particularly low marginal costs, even in comparison to heroin. The other side of the coin is that fentanyls are so potent that using them is dangerous especially without the skills and equipment for accurate dosing.[16] As market participants acquire these skills, or fentanyls are produced more cheaply, opioid consumption could further jump up in the way illustrated in Figure 1b.

III. A Database of Federal Opioid Policies

Although dozens of public policies can affect opioid markets, the sufficient statistics framework suggests categorizing them according to their effect on the structure (Rx vs Im) of opioid supply and demand. One category of policies affects the marginal and average costs of illicitly-manufactured opioids to final consumers. A second policy category affects the marginal and average costs at various points in the supply chain for prescription opioids. Because the magnitude of the jump terms in equations (2) - (4) increase with the level of the prescription price, this category can be further subdivided in terms of the price point most affected. The (illicit) secondary prescription market has the highest price point, followed by cash purchases at pharmacies, followed by pharmacy purchases covered by insurance. The third policy category shifts opioid demand in the sense that the policies change the prices or availability of opioid substitutes and complements such as addiction treatment, prescription tranquilizers, or overdose medications. Finally, each policy should be shown in historical context with the others, and with evidence of their behavioral effects, because the sign and magnitude of one policy's effects depends on the status of the others.

[16] Fentanyl itself has been the "most popular opioid analgesics in modern anesthesia and pain therapy" because of its "potency, speed of onset, and relatively short duration" (Stanley, Fentanyl 2005). Per gram, fentanyl is at least 15 times more potent than heroin. Carfentanyl is at least a hundred times more potent than that.

Table 1 shows the results of this exercise for federal policies during the years 2000-19 and serves as a point of reference for the price and behavioral measures presented in the rest of this paper. Policies for analysis were identified from the Federal Register and agency press releases.[17] For measurement purposes, the table also categorizes Rx incentives more finely according to the chain of Rx production, which includes prescribing and consumer effort and expenditure, among others. The table reveals a few patterns. First, opioid prescribing was subsidized and saw restrictions eased through about 2012, with a partial reversal in the years thereafter. Second, subsidies to patients began in 2001 and continue at least through the end of the sample period. Patients eventually experienced the effects of tighter regulation, especially in 2011 when a primary Rx opioid brand was replaced with a new "abuse-deterrent formulation." In Table 1's law enforcement column, we see steps in both directions in terms of the "War on Drugs" generally, and fentanyl specifically. A couple of policies relate to opioid substitutes and complements. These patterns are discussed more extensively in what follows. A longer version of this paper (Mulligan, 2020) provides additional details on the less significant prescriber regulations shown in Table 1.

III.A. Prescriber Subsidies and Regulations

In 2000, the Veterans Health Administration (VHA) mandated pain as "the 5th Vital Sign," which meant that pain would be routinely screened and documented with the other vital signs. Moreover, the patient's other vital signs and behavior "should not be used instead of self-report" by the healthcare provider making the pain assessment (Department of Veterans Affairs, 2000).[18] The new standards would soon come into civilian practice too, with three phases of financial encouragement from the Centers for Medicare and Medicaid Services (CMS).[19] CMS has long conditioned hospital reimbursement under its Medicare and Medicaid programs on

[17] Appendix II provides more detail on the search criteria and algorithm.
[18] An exception was allowed for patients unable to communicate. At its peak in 2012, the VHA would be dispensing opioid prescriptions to almost 700,000 unique patients per quarter (Good 2017).
[19] Centers for Medicare and Medicaid Services (2017) and Pacula and Powell (2018). CMS is the component of the Department of Health and Human Services that administers the two major federal health insurance programs (as well as the smaller although disproportionately famous "Obamacare" programs).

adequate hospital quality, to be assessed by the Joint Commission on Accreditation of Healthcare Organizations, another accreditation organization, or state survey agencies.[20] In 2001, the Joint Commission made pain management part of the accreditation process, beginning the first phase of financial encouragement by CMS.[21]

The second major financial incentive began in 2007 when CMS would withhold two percent of full reimbursement if a hospital failed to participate in a patient survey. The survey, known as the HCAHPS survey, included questions about the patient's pain management experience.[22] As described by Physicians for Responsible Opioid Prescribing (2016), physicians learned that prescribing extra, or more potent, painkillers tended to produce higher HCAHPS scores, and therefore additional funding, for their hospital. This CMS incentive remained for 12 years until the pain treatment questions were removed from the survey in October 2019.

The third incentive phase began in 2012, when CMS implemented the "valued-based purchasing" requirement from the Affordable Care Act. Pursuant to the law, "value-based incentive payments to hospitals were tied to the value of these patient experience performance measures, which included pain management scores as a core component" (Pacula & Powell, 2018). Mulligan (2020) estimates that, as of 2016, the roughly $5 billion annual prescription opioid market was subsidized about $0.7 billion annually from the combination of these three financial incentives.

In 2017 the President's Commission on Combating Drug Addiction and the Opioid Crisis (2017) concluded that CMS should end its financial incentives for over-prescribing opioids. CMS followed the Commission's recommendation, removing the pain-communication questions from HCAHPS effective October 2019 (Centers for Medicare and Medicaid Services, 2019). The 2018 SUPPORT for Patients and Communities Act included provisions for tightening prescriber guidelines and restricting the import of illicit drugs.

[20] CMS is required "to ensure that the Joint Commission's surveying of accredited hospitals is equivalent to state agency surveying of unaccredited hospitals" (Lohr 1990, p. 131).
[21] At the same time, HHS delegated the regulation of opioids for drug treatment programs to JCAHO, including removing HHS caps on dosage and days supply (66 FR 4076) and later expanding the list of approved opioids (68 FR 27937).
[22] The November 2006 rule promulgating this requirement was 71 FR 68193. The two percent penalty come from Section 5001(a) of the 2005 Deficit Reduction Act.

Earlier, VHA began its Opioid Safety Initiative in 2013, which would prove to be the beginning of a significant (at least 25 percent) decline in the number of unique VA patients dispensed an Rx opioid (Good, 2017). The initiative included urine screening for opioid abuse, prescribing with tapering protocols, offering substitute treatments for chronic pain, and using state-level prescription-drug monitoring programs (United States Government Accountability Office, 2018). To the extent that they deplete supply to secondary markets, "Prescription Take-back" programs are also policies that disproportionately affect higher-priced Rx opioids and therefore more likely to have the unconventional effect of increasing total opioid consumption.

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 misusers (Food and Drug Administration, 2018). Because the reformulation is especially revealing as to the number of consumers on the margin between Rx and Im, its discussion is deferred until Section V.

III.B. Patient Subsidies

Consumers with prescription-drug coverage have a lower out-of-pocket price for Rx opioids than cash customers do. As federal programs such as Medicare or Medicaid expand, then more consumers face the lower out-of-pocket price and federal taxpayers pay most of the difference. Of the various federal coverage expansions since 2000, the largest from the perspective of the opioid market was the launch of Medicare Part D in 2006. Prior to that date, Medicare, which is the federal health insurance program for the elderly and disabled, did not cover retail prescription drugs. Within a year, Part D enrollment had exceeded 30 million, cutting their Rx opioid out-of-pocket price by more than 90 percent compared with no coverage.[23] Council of Economic Advisers (April 2019, p. Figure 12) shows that the largest single-year increase in opioid prescriptions occurred in 2007.

[23] 2007 enrollment is from Table IV.B7 of Boards of Trustees, Federal Hospital Insurance and Federal Supplementantal Medicare Insurance Trust Funds (2014). The average out-of-pocket opioid price for Medicare is from Council of Economic Advisers (April 2019, p. 6) and refers to Oxycontin.

Many of the Rx opioids paid by Medicare were consumed by individuals ineligible for Medicare. Some of those prescriptions were given away to friends or relatives. Others were stolen or sold. The diversion of Medicare prescriptions was significant enough that Powell, Pacula, and Taylor (2017) found that the opioid-related fatality rate among persons *ineligible* for Medicare increased significantly as a result of the creation of Medicare Part D.

III.C. Policies toward Opioid Substitutes and Complements

The federal document search revealed only a couple of regulations related to opioid substitutes and complements, which are included in Table 1 without making a dedicated column. One of them is Section 2502 of the 2010 Affordable Care Act (ACA) that requires all health plans to cover benzodiazepines (benzos), which are a complement to opioid consumption, and thereby reduce their out-of-pocket cost. This overturned the prior practice, especially in Medicare, of specifically excluding benzos from coverage (Bambauer, Sabin, & Soumerai, 2005; Centers for Medicare and Medicaid Services, 2005; 2012).

Benzos are quantitatively important in opioid markets because they are part of the opioid-benzo cocktail that is a favorite among opioid misusers. The tranquilizers "enhance" the feeling of opioid consumption, regardless of whether the opioids themselves are sourced from Rx or Im. For the same reason, benzos carry serious risk of death when used in combination with opioids (Sun, et al., 2017). A study of 2,400 veterans who died from a drug overdose found that 49 percent of them had been prescribed concurrent benzos.[24] According to the MCOD files described further below, about a quarter of prescription opioid overdose fatalities in the general population have involved benzos. The annual number of fatalities involving both benzos and opioids appear to have peaked in 2017 at almost 10,000.

The 2018 Support Act also included provisions reducing regulatory barriers for treating opioid-abuse disorder, encouraging the distribution of naloxone (a drug that can reverse opioid

[24] Park et al. (2015). The primarily male study sample was selected as the 112,069 U.S. veterans who received VHA prescriptions for opioids during 2004-9. Of those, 2,400 died from drug overdose during the study time frame. Also note that, prior to 2017, the FDA did not require benzo prescriptions to carry a warning about the potentially lethal opioid interactions (Food and Drug Administration 2016).

overdose), and increased funding for those treatments (Food and Drug Administration, 2019b). The Comprehensive Addiction and Recovery Act of 2016 also included provisions in some of these general areas.

IV. Price and Quantity Data

With the exception of the calibration exercise at the end, the key quantitative propositions about measured prices in this paper are that: (i) heroin was significantly more expensive per MGE than Rx opioids in the 1990s, (ii) more recently, especially when mixed with fentanyl as it often is in actual markets, heroin is significantly cheaper, and (iii) a significant drop in the relative quality-adjusted price of heroin occurred since 2013. Each one of these price propositions, as well as various qualitative hypotheses about quantities, requires data from illicit markets where market participants have strong incentives to avoid being measured. Three precautions are taken in this paper to avoid conclusions contaminated by measurement errors and source limitations. The first is to advance propositions that are expected to be robust to measurement error, e.g., "heroin was significantly more expensive" rather than "heroin cost 43 percent more." Second, especially regarding quantities, I employ multiple independently-sourced indicators of the same behavior such as fatality rates from medical examiners, results from household surveys, and reports from law enforcement (Pardo, 2019). This approach is sometimes possible with price measures too, by comparing retail transaction prices reported by law enforcement with household indicators and with wholesale cost information from law enforcement. The third precaution, following Arkes et al (2008) and others is to view the price findings in the context of the quantity findings.

IV.A. The Structure of Opioid Prices

Market participants have described a situation in which the per-dose price gap between heroin and prescription opioids has 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" (Butterfield, 2001; Jayawant & Balkrishnan, 2005; Quinones, 2015). Years later, heroin is recognized to be the cheaper alternative (Cicero, Ellis, Surratt, & Kurtz, 2014; Cicero, Ellis, & Kasper, 2017; National Academies of Sciences, Medicine, & others, 2017). A recent survey of people in treatment for opioid addiction found that “almost all—94 percent—said they chose to use heroin because prescription opioids were ‘far more expensive and harder to obtain’” (National Institute on Drug Abuse, 2018a).

Quantifying prices on the prescription side of the gap is more traditional because prescriptions are manufactured in licit markets. Prescription sales and invoice price time series for 1992-2016 are from FDA (2018), which it compiled from IQVIA National Sales Perspectives. Prescription out-of-pocket prices are from Council of Economic Advisers (April 2019), which it compiled from the Medical Expenditure Panel Survey for 2001-2015. Heroin prices are from United Nations Office on Drugs and Crime (2018) and the various Drug Threat Assessments published by the U.S. Drug Enforcement Administration (DEA). These primarily reflect heroin prices observed by DEA as a byproduct of its efforts to prosecute offenders and to monitor the drug supply, rather than for research purposes. DEA samples, which since 2007 are not publicly available at the transaction level, are unlikely to be representative and need cleaning (Reuter & Caulkins, 2004). On the other hand, as shown below, the overall downward trend is large and consistent with known cost drivers.

Since 2013, consumers frequently receive heroin mixed – some would say adulterated – with fentanyl. Unlike other traditional heroin adulterants or fillers, fentanyl produces morphine symptoms more powerfully than the heroin itself. As a result, heroin prices per pure heroin MGE are above prices per MGE contained in retail heroin products, which are not part of DEA’s price and purity data. The gap between DEA prices and price per MGE is significant because fentanyl “is phenomenally inexpensive per dose in wholesale markets” (Pardo, 2019, p. p. 119). Fentanyl is enough cheaper than pure heroin to largely displace heroin from illicit markets, as it has done in some countries and regions of North America.[25]

The introduction of fentanyl is expected to significantly reduce Im prices per MGE because, on the scale of the markups and costs of transit and delivery along the illicit supply

[25] See Pardo (2019, pp. 20, 109, 125) on British Columbia, Estonia, and perhaps Latvia. Mortality and NFLIS data suggest that this had also occurred in most of the northeast U.S. by 2019.

chain, fentanyl essentially eliminates the costs of manufacturing the opioid ingredient.[26] Moreover, given its small weight and volume per MGE, fentanyl is also cheaper to transport and deliver to the final consumer. Quantifying the amount of savings is difficult, but suppose for the moment that costs are passed one-for-one into retail price and that fentanyl can travel the domestic part of the production stream at the same $750 cost per pure gram as heroin. That puts retail prices at essentially $50 for the morphine equivalent of a gram of heroin as compared to $800 for heroin itself.[27] As other ingredients are added to the weight and volume of the fentanyl products delivered at retail, costs are added to the $50 but the magnitude of these additions falls as they are incurred closer to the final consumer.[28] At the same time, changes in law enforcement that reduced transit, delivery, and retailing costs for all illicit opioids may have disproportionately affected fentanyl. To my knowledge, the only times series on the role of fentanyl in illicit-opioid prices is the one published by the Council of Economic Advisers (April 2019), which I update and use to adjust the DEA price per heroin gram.[29] Given the data-source limitations, I also show sensitivity analysis.

Using a logarithmic scale, Figure 2 shows results for the relative out-of-pocket price of the two main opioid alternatives: illegal heroin (including any adulteration with fentanyl) and prescription (Rx) opioids. As explained further in Appendix III, the former is from DEA reports, adjusted using CEA's series to reflect price per MGE; the adjustment is negligible until 2014 when fentanyl emerges on a large scale. The Rx series used in the figure is an index of branded and generic invoice prices from FDA, which does not reflect third party-payments that would lower the price or the premium for diverted prescriptions that would raise the price (or, for a person purchasing from a pharmacy, the opportunity cost). Figure 2 suggests that, between 2005 and 2013, a given expenditure bought, on average, only slightly more morphine equivalents of

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[26] A kilogram of pure fentanyl for delivery to the U.S. can be purchased from Chinese sellers for $2,000 to $5,000 (Pardo, Davis and Moore 2019). The morphine equivalent amount of pure heroin (15 kg) would cost $750,000 on the Mexican side of the border.

[27] Using the data from the previous footnote and 15-to-1 relative potency for fentanyl, $50 \approx (800 - 750000/15000 + 5000/1000)/15$.

[28] Harm Reduction Ohio (2019) finds that fentanyl is mixed with other drugs only after the drugs arrive in local communities and are subdivided in small portions. By all accounts, fentanyl is more dangerous than heroin per MME, at least in the short run until market participants acquire the tools and expertise to dose accurately.

[29] But see also Dismukes (2018, p. 183) who finds that, in one-gram quantities, by 2017 heroin cut with fentanyl was about half the price of heroin by itself.

prescription opioids. In contrast, a given expenditure bought five times as much Rx in 1992 whereas in 2017 it bought only one third. This, together with the fact that substantial numbers of consumers have been on the margin between Rx and Im, raises the concern that price effects on opioid consumption are quite different in, say, 2010 than they were in 1992 or in 2017.

Heterogeneity of Rx opioid sources is relevant for policies that target specific segments of the Rx market, such as the 2010 FDA approval of a reformulation of a leading opioid brand. A longer version of this paper assembles a cross-section of prescription opioid prices in 2014 and 2015 (Mulligan, 2020), finding that the average branded price per morphine-gram equivalent (MGE) was, net of the average insurance third-party payment, $64 when purchased at a pharmacy but $880 on the secondary market (Schnell, 2018). The quantity-weighted average across sources was $336 per MGE.

IV.B. Opioid Fatalities and other Indicators of Opioid Consumption

Fatalities are measured from the Multiple Cause of Death Files (MCOD) 1999-2019 (National Center for Health Statistics, 1999-2019), which contain a record for every death certificate filed with a U.S. state or District of Columbia (essentially every death in the country). Each record shows an underlying cause of death as well as up to 20 contributing causes of death (Redelings, Wise, & Sorvillo, 2007). I select only those records where the underlying cause of death is drug poisoning, which are 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 in which 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).[30] Except where otherwise noted, I take opium, heroin, and synthetic as illicitly manufactured and the other two T codes as Rx opioids. I also present sensitivity analysis to including death certificates with drug poisoning from benzodiazepines (tranquilizers T code 42.4), cocaine (40.5), antidepressants (43.0-43.1), psychostimulants (43.6, especially methamphetamine), and unspecified drugs (40.6 and 50.9).

[30] Although autopsies are generally uncommon, they occur in more than 80 percent of drug deaths. The autopsy rates are even greater for the non-elderly and for deaths that occur outside a medical facility. Conditional on age, sex, race, year, and place of death, autopsy rates for the different T codes are within a half percentage point of each other.

For the purpose of measuring fatality rates by demographic group or for the nation, I use population estimates from the National Cancer Institute's Surveillance Epidemiology and End Results (SEER) program by sex, year, integer age and white/black/other. 2019 fatality rates use the most recent SEER year (2018) as their denominators.

Where proxies for heroin consumption are needed that are distinct from death rates, I use the National Survey on Drug Use and Health (NSDUH) for the years 2002-19, which covers ages 12 and older.[31] Respondents are asked whether they use various specific drugs, including heroin and nonmedical use of prescription pain relievers but not fentanyl. 1.5 percent of the survey has ever used heroin. 0.3 percent used heroin in the year prior to the survey, with about half of those using in the past month. Figure 3's black series displays the prior-month heroin series, displayed as an index. The prevalence almost doubles between 2010 and 2016.[32]

The National Forensic Laboratory Information System (NFLIS) provides a very different way of measuring the consumption of drugs obtained from criminal enterprises. It receives about 1.5 million reports annually from 279 forensic laboratories as to drugs found in "substances secured in law enforcement operations" (Drug Enforcement Administration, Diversion Control Division, 2020, p. p. 4). These laboratories handled 98 percent of the national drug caseload in 2019.[33] The majority of the drug reports are cannabis/THC, methamphetamine, or cocaine. Opioids range from less than one percent to 13 percent of drug reports, depending on the year. As indicated by their sheer numbers, these reports are not directly tied to fatalities although fatalities in the community may affect the allocation of law enforcement effort among various types of drug criminals.

Figure 3's red series shows heroin's share of the reports. The similar time pattern for the two series suggests that NSDUH might reflect genuine changes over time even though the survey significantly underestimates the prevalence of drug abuse (Caulkins, Kilmer, Reuter, & Midgette, 2015).

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[31] Because the survey was redesigned in 2015, I do not make comparisons across 2014-15 without specifically noting how a variable seems to be comparable over that time.

[32] See also DEA (2016, p. 58) citing "a large influx of new [heroin] users."

[33] It is the laboratory's decision whether to conduct testing, which they often do not when "the criminal case was dismissed from court or if no defendant could be linked to the case" (Drug Enforcement Administration, Diversion Control Division 2020, p. 29).

NFLIS also helps quantify the dynamics of aggregate consumption of illicit synthetic opioids (overwhelmingly, illicit fentanyl), especially when supplemented with two additional types of information. The first addition is an estimate of the number of fentanyl MGEs acquired in the average law enforcement operation as compared to the average for heroin. My estimate is from U.S. Customs and Border Patrol, assuming that their seizures are representative in terms of the ratio of the average fentanyl seizure to the average heroin seizure. Second, and less significant, is an estimate of the comparatively small amount of misuse of synthetic-opioid prescriptions (as distinct from illicitly manufactured synthetic opioids such as illicit fentanyl) that are thought to be responsible for the deaths in the early 2000s coded more coarsely as synthetic-opioid overdoses. Figure 4's red series shows the result, assuming for the sake of illustration that deaths per MGE consumed are the same for heroin and synthetic opioids.[34] Also shown are two ingredients of the red series: the death rate from prescription opioids and, especially, the death rate from heroin.

The timing and, for many of the changes, magnitudes are similar between Figure 4's red series and actual deaths from synthetic opioids (black), despite the fact that the non-NFLIS component series are non-monotonic after 2010 while the actual deaths increase at an increasing rate. The similarities suggest that either fentanyl deaths track fentanyl consumption or that their departures happen to be largely offset in errors in measuring fentanyl consumption from NFLIS. At minimum, the NFLIS data is evidence that a large fentanyl-driven increase in MMEs consumed has coincided with a large increase in deaths from fentanyl, even if the latter is also partly driven by additional dangers of consuming a given number of MMEs in fentanyl form. Figures 3 and 4 together show that both heroin fatalities and heroin consumption grew significantly, especially following the year 2010. Heroin fatalities and consumption peaked around 2016 while fentanyl fatalities and consumption were increasing sharply.

Aggregate data show reductions in opioid prescriptions after 2010, and significant reductions after 2012 when multiple policies were encouraging consumers to shift from Rx to Im (Table 1). FDA (2018) data show aggregate morphine equivalent opioid prescriptions in 2011

[34] The red series is the sum of a term representing synthetic-opioid prescriptions, estimated as the MCOD fatality rate T40.2 time series rescaled by the 2001 ratio of T40.4 to T40.2, and a term representing illicit synthetic opioids estimated each year as the product of MCOD heroin fatality rate times the ratio of fentanyl MGEs to heroin in NFLIS. The first term is always less than 1.1 deaths per 100,000 population.

were about one percent below their 2010 peak.[35] In 2012, they were about three percent below the peak. By 2016, Rx opioids were 20 percent below their peak. Administrative data tell a similar story as to the number of prescriptions. Medicare opioid prescriptions per Part D enrollee fell 16 percent between 2013 (the peak year) and 2017 (Buchmueller & Carey, 2018; Centers for Medicare and Medicaid Services, 2019b). Medicaid opioid prescriptions fell 4 percent from 2015 to 2016, and another 14 percent the next year (Centers for Medicare and Medicaid Services, 2019a). Express Scripts (2019), which has about one quarter of the pharmacy benefit management market (Paavola, 2019), reported that opioid prescriptions fell 17 percent from 2017 to 2018 for commercial plans, 13 percent for Medicare plans, 25 percent for Medicaid plans, and 19 percent for plans on the health exchanges.

V. Evidence that Consumers were on the Margin of Rx and Im

The theoretical framework suggests that Rx and Im markets could be connected in three ways. Mixed Rx-Im consumption is one connection, because by definition such a consumer is participating in both markets. Thirty-eight percent of overdose deaths in 2018 involving Rx opioids also involved Im opioids. Only five years earlier (2012), the percentage was only 12. This is suggestive of simultaneous participation in both markets during some of the more recent years, although the mortality data does not distinguish Rx chemicals that the consumer obtained through licit channels from Rx chemicals obtained and administered in essentially the same way as heroin.[36]

The second connection is over the life cycle, with perhaps younger people beginning with prescription opioids and then transitioning to illicit markets as they consume increased amounts or experience reductions in their fixed cost. The third connection is in the cohort dimension, where older cohorts started on Rx opioids whereas a significant fraction of later cohorts (or different race/ethnic groups) began opioid misuse with Im opioids because they face a different

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[35] Propoxyphene + Acetaminophen, an opioid heavily prescribed for mild to moderate pain, was withdrawn from the market at the end of 2010 pursuant to a FDA recommendation (Food and Drug Administration 2010). DEA classified it as Schedule IV "with a low potential for abuse and low risk of dependence." For the purposes of this paper, it does not matter whether or how much this withdrawal versus OxyContin reformulation caused the shift from Rx to Im.
[36] Im-Im combinations are more commonly reported in the MCOD records than Rx-Im combinations. See also Warner et al (2016).

set of prices than the early cohorts did. The evidence that follows suggests that both of these connections are empirically important at various points during the years 2000-2019, which means that the CROSS elasticity featured in equation (3) is rarely well approximated by zero. Section VII looks at these results in the context of equation (3) by also calibrating the other sufficient statistics.

V.A. Opioid Sourcing among Adolescents and Young Adults

The National Institute on Drug Abuse reports that a "study of young, urban injection drug users interviewed in 2008 and 2009 found that 86 percent had used [Rx] opioid pain relievers nonmedically prior to using heroin."[37] Although initiation patterns may be different after 2014 when Im prices fell relative to Rx (Figure 2), the NIDA survey by itself indicates that it has been common for individuals to begin opioid misuse with prescriptions and then transition to Im opioids. A broader perspective is shown by Figure 5's shares of opioid fatalities that involve Rx but not Im opioids, as measured in the 1999-2019 MCOD files, by age and time period.[38] Year effects, which capture market prices and other time-specific factors, are held constant by means of least-squares regression in the sample of deaths at age 14-25. In the earlier period (blue series), deaths involving Im opioids were relatively rare at ages 14 and 15. They became more common with age, although even at age 25 Im opioids were involved (and recorded in death certificates) in a minority of opioid deaths. More recently (red series), the Rx share is lower at every age although the age pattern of Rx share changed little over time. During both periods, the per-capita fatality rates indicated by marker size continue to increase with age after transition from Rx to Im, suggesting that consumption may follow a similar pattern.

Figure 5 suggests that regulatory efforts to reduce opioid prescribing that began in 2011 (recall Table 1) might have reduced opioid fatalities the most among adolescents due to the

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[37] National Institute on Drug Abuse (2018a). It further notes that, in contrast, "of people entering treatment for heroin addiction who began abusing opioids in the 1960s, more than 80 percent started with heroin." The frequency of initiation to opioid abuse via Im opioids may be increasing in recent years (National Academies of Sciences, Medicine and others 2017, Chapter 4).

[38] Only ages 14-25 are shown because those are the typical ages of initiation to heroin and prescription-opioid misuse, and the shares change little after age 25.

apparent Rx intensity of their opioid use at the same time causing some adolescents to use Im opioids instead. This hypothesis is further supported in Figure 6, which shows overall opioid fatality series (Rx and Im) by age. Ages 14-17 (adolescent), 18-21 (young adult), and all ages follow quite similar time patterns through about 2010. Thereafter gaps emerge between the series. The fact that the adolescent series is the only one to remain below its 2010 values suggests that, against falling Im prices and other factors increasing fatalities, tighter prescription policies may have reduced fatalities in adolescents but not enough to last into young adulthood when the Rx share of opioid consumption would have been low regardless (Figure 5). In the context of equation (3), the average CROSS and average $1-r$ among adolescents may be less than they are at older ages because the adolescents have a higher ratio of fixed cost to the quantity that they would consume if sourcing from Rx. Their total opioid demand is therefore more likely than the others to slope the "right" way with respect to the Rx price.

Beyond adolescence, the age and sex patterns are different. Figure 7 displays fatality rate changes by age and sex after 2010, which is a timeframe coinciding with policy changes discouraging Rx consumption and encouraging Im consumption (more on this below). Most of the groups show, to varying degrees, reduced fatality rates from Rx opioids alone. The economically and statistically significant relationship with the change in the fatality rate involving any opioids is negative, which is remarkable given that the Rx fatalities measured on the horizontal axis are included as part of the any-opioid fatalities measured on the vertical axis. Using the 48 single-year age groups 18-65, the OLS regression line of the change in the any-opioids rate on the change in the Rx-only rate has a slope of −1.8 (p-value < 0.001), which means that a reduction in the Rx-only fatality rate (per 100,000 population) by one is associated with 1.8 *more* opioid fatalities. Consistent with the hypothesis that many opioid misusers were near the margin between Rx and Im, the Rx-only reductions shown in Figure 7 are associated with the demographic groups for whom we might suspect comparatively low fixed costs of illicit activity: men and persons in their 20s and 30s.[39]

[39] If Rx-only fatality rate changes are predicted in a first stage regression on sex and an age quadratic, all three of the coefficients are economically and statistically significant. The first stage adjusted R-squared is 0.75. The TSLS estimate of the coefficient on Rx-only in the all-opioids equation is −2.8 (p-value < 0.001).

V.B. The Reformulation of Oxycontin in 2010

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. Schnell's (2018) analysis of IMS data show that prescriptions of OxyContin (the primary Rx opioid that was reformulated at that time) fell sharply later in the year.[40] By January 2011, OxyContin prescriptions fell below 2006 levels. Several studies of the effects of reformulation have been conducted, which can now be interpreted and placed into a broader context by using my sufficient statistics results.

Recall from Figure 2 and Table 1 that 2010 and 2011 are before cheap fentanyl had so deeply penetrated illicit markets. In terms of equation (3), p_R/p_I was closer to one, and not changing much, at the time of the reformulation than it would be 5 years later. All else the same, reformulation is expected to initially increase total opioid consumption less, or reduce it more, than would a Rx price increase that occurred later. Ruhm (2019) 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.[41] It follows from their results that equation (3) evaluates to about zero in the years studied (2011 through about 2013) and therefore would evaluate to a positive number more recently when p_R/p_I is closer to one. A more quantitative statement of this result is shown in Section VII.

Recall from Figure 3 that both the NSDUH and NFLIS proxies for heroin consumption increase at a higher rate after 2010, which is the same year that Evans, Lieber and Power (2019)

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[40] As of 2019, no generic abuse-deterrent opioid painkillers had yet been approved by the FDA (Food and Drug Administration 2019a).

[41] Note that 2013 and 2014 are the most recent years used by Alpert Powel and Pacula (2018) and Evans, Leiber and Power (2019), respectively, to measure the aggregate mortality effect of the reformulation. Schnell (2018) also finds that reformulation of Oxycontin increased sales of other Rx opioids.

find a structural break in the heroin fatality rate (see also my Figure 4's circle series). This occurred before cheap fentanyl became widely available. In percentage terms, the fatality rate increases more (and falls more after 2015 or so). The quantitative differences may indicate switching by particularly high-volume Rx consumers, that heroin consumption was more dangerous for the new heroin users, or that the consumption series have measurement errors that are correlated with actual consumption.[42]

V.C. Did Access to Prescription Opioids Mitigate the Effect of Fentanyl on Fatality Rates Among Whites?

Age patterns such as those shown in Figure 5 suggests that Rx can be a gateway to Im; the two are complements over time for some purposes. However, while fully consistent with habit formation, the theory also allows them to be substitutes. A population that faces high fixed and marginal costs of Rx could be more affected by a reduction in Im prices because the Rx market is unattractive to them. African-Americans may be such a population. They have been more likely than whites to be uninsured (especially among young men and less so among older women), and thereby have to pay more out of pocket for prescriptions, both at the pharmacy and for a physician-office visit to obtain a prescription.[43] Some studies have found that, referring to prescription opioids, "racial/ethnic minorities consistently receive less adequate treatment for acute and chronic pain" (Mossey, 2011). Swift et al (2019) conclude that racial discrimination is a factor preventing prescription opioid use among blacks. Regardless of the mechanisms, the MCOD data show that by 2010 the fatality rate involving Rx opioids among whites was triple than among blacks. Among blacks, the age gradient of the Rx share of opioid fatalities is half of what is shown in Figure 5 for the entire population and statistically indistinguishable from zero at the conventional significance levels.

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[42] A pattern of opioid addiction originating with prescription followed by a switch to more potent heroin when prescribing practices were tightened has historical precedents. Writing before the most recent episode became apparent, Cartwright (2001, Chapter 4) describes the same pattern early in the 20th century.
[43] The NSDUH data used in this paper, among others, show racial disparities in health insurance.

Section II's theory predicts that a reduction in p_I would increase the black-white gap in opioid fatalities (if the gap is negative in the baseline, then it would be less negative). In terms of Figures 1a and 1b, the change for whites resembles B to C whereas the change for blacks resembles A to C. Without necessarily changing that gap, an increase in p_R would reduce the black-white gap in Im consumption. These price effects on the racial composition of total opioid fatalities would, in theory, be particularly pronounced in the age-sex groups that begin with larger race gaps in the prescription-intensity of their fatalities. Figure 8 shows the black-white gap in fatalities from drug overdose for various age-sex groups. Young men and older women are selected because they are the age-sex groups with the more extreme race gap for Rx fatalities (not shown in the figure): especially large for the former and essentially zero for the latter. Between 2010 and 2013, which were the first year of the reformulated prescription opioids, there was little change in the race gap overall, for young men, or for older women. Nevertheless, deaths involving heroin, which was the primary form of Im opioids at the time, increased significantly more for whites than for blacks, especially for young men.

For many years opioid fatalities were more common among whites, which was viewed as somewhat of a puzzle (Case & Deaton, 2020, p. Chapter 5).[44] More research is needed on similarities and differences between the races in terms of addiction to opioids and other drugs, but the theory and Figure 8's data suggest that part of the answer may involve the distinction between prescription sources versus illicit markets (see also Alexander, Kiang and Barbieri 2018). During the period 2001-2010, prescription opioids became cheaper whereas later, especially 2014-2019, cheap fentanyl increasingly became available. During the earlier period, opioid fatalities increase more for whites than for blacks, while the opposite happened later. The 2014-19 changes were large enough that the national black rate exceeded the white rate for the first time in the 21st century.

[44] The "All age & sex" series in Figure 8 would not change much if fatalities involving cocaine, antidepressants, benzodiazepines, and methamphetamine were also included. In particular, the white rate would be significantly above the black rate for most of the years 2001-18, but in 2019 the black rate would be the greater of the two.

VI. Evidence of Illicit Supply Shifts beginning in 2013

The supply of illicit opioids appears to shift beginning in 2013, three years after OxyContin's reformulation. The section uses the price and quantity measures to confirm that, although high opioid demand may have been a factor, a supply shift associated with the distribution of illicit fentanyl was large by historical standards. Technological progress and law enforcement changes are shown to be potentially important cost shifters during this period.

VI.A. Quantity and Price Series Turn

The year 2013 was a turning point for opioid fatality rates (Rx and Im combined), fentanyl and heroin in crime labs, heroin usage, quality-adjusted heroin prices, and survey reports of ease of heroin access. One way to quantify the trend change is, following Evans, Lieber, and Power (2019), to estimate a structural break. The quarterly time series for unintentional drug poisonings involving opioids per one hundred thousand of population aged 16 and older between 2005 and 2017 has its break in 2013-Q3.[45] However, for several of the quantity series the changes are extreme enough that the pre-2013 distribution of changes has little or no overlap with the post-2013 distribution. For example, Figure 4 shows how fatalities involving synthetic opioids and (indirectly) drug reports of fentanyl increased more than ten (per 100K population) between 2013 and 2019, after increasing less than one in the prior dozen years. Figure 9 assembles four such annual series and highlights the right-tail outliers in their changes. Here the right tail is defined to be the top quartile of changes. Each outlier's magnitude is measured as its distance above the 75th percentile, with distance normalized so that the entire series of changes has interquartile range of one. Figure 9 shows only one series with any right-tail outliers before 2014, which is the NSDUH series for monthly heroin usage. The years 2014, 2015, and 2017 each have three right-tail outliers out of a maximum possible four. All four series have an outlier in 2016. Right-tail outliers are less common in each of 2018 and 2019.

[45] The date is estimated using the supremum Wald test (Perron 2006), as implemented by STATA. The Wald statistics are similar for alternative structural breaks one quarter early, one quarterly later, and two quarters later. Structural break analysis generally can be sensitive to the sample period.

This well-known increase in the trend for opioid fatalities are not merely the delivery of a constant quantity of Im opioid consumption in more dangerous form. As illustrated by the red series in Figure 4 and the blue series in Figure 9, the law enforcement data point to such a large increase in the amount of fentanyl in the country that the total MMEs supplied must far exceed what it was in 2013 and the years before. Figure 3, derived from both law enforcement data and household survey data (NSDUH), suggests that heroin quantities also increased between 2013 and 2015 and perhaps two years beyond. The acceleration of heroin usage between 2013 and 2014 is notable given that three years earlier the reformulation of the leading Rx opioid brand also encouraged Rx consumers to switch to heroin.

Council of Economic Advisers (April 2019) reports little change between 2008 and 2012 in the inflation-adjusted price (per MGE) of Im opioids. It estimates that prices were nine percent lower in 2013 (not noticeably outside the 2008-12 range), followed by drops of 22 percent and 34 percent in the next two years. Although NSDUH does not measure prices, it does ask respondents (including those who do not use heroin) whether heroin is easy to obtain. The fraction responding affirmatively has been between 15 and 20 percent, with the two largest annual increases occurring in 2013-14 and 2015-16, with no close third place.[46]

VI.B. Fentanyl in Historical Context

Historically pharmaceutical fentanyl has, on a small scale, been diverted from healthcare facilities and pharmacies for non-medical use (Stanley, 2014; Drug Enforcement Administration, 2016, p. p. vii). This paper refers to the much larger quantities of non-medical fentanyl that originate from illicit manufacturing facilities, especially in Mexico, China, and India. My search of Department of Justice press releases between 2001 and 2019 reveals several episodes of fentanyl entering the U.S. heroin supply before law enforcement could shut it down.[47] "From 1990 through 2005 at least nine clandestine fentanyl laboratories were seized in the United States" (U.S. Department of Justice, 2006). Fentanyl was illicitly manufactured in Mexico for the U.S. market in 2006, to the point where the Office of National Drug Control Policy

[46] 2014-15 is unavailable due to a change in NSDUH survey design.
[47] Ten other illicit fentanyls episodes have been cited for the period 1979-88 (Polisen 2018, p. 9).

concluded that "drug traffickers have substantially poisoned the drug supply in the United States" by adding fentanyl to heroin.[48] Law enforcement shut down that fentanyl manufacturing, although not before national fatality rates involving synthetic opioids briefly spiked 72 percent above their previous levels, with deaths concentrated in Illinois, Missouri, Michigan, Ohio, and West Virginia.[49]

Both the death records and NFLIS data suggest that illicit fentanyl re-entered U.S. markets in about December 2013 (Drug Enforcement Administration, Diversion Control Division, 2017). Fentanyls of Chinese origin were introduced in Sweden in January 2014 (Polisen, 2018, p. p. 15). As of 2018, the fentanyl analogs were reportedly no longer available on the Swedish market due to arrests and prosecution of their vendors (Pardo, 2019, p. p. 92).

Perhaps unaware of how illicit markets would evolve, parts of government in the U.S. were less willing to maintain or escalate the war on drugs. As Eric Holder describes it, he and then Senator Barack Obama were mutually aware of the costs imposed on low-level drug offenders by federal sentencing rules. They agreed on "the need for change, a need for new approaches" and seized their chance to make such changes when Obama became president and made Holder his attorney general (Breslow, 2016). In August 2013, Holder issued his famous "Holder Memo" directing federal lawyers to stop prosecuting nonviolent drug crimes (Holder, 2013). The new initiative, which he would credit with ending 33 consecutive years of increases in the federal prison population, would be called "Smart on Crime" and "nothing less than historic."[50]

VI.C. Determinants of Cost in Illicit Markets

It goes without saying that prison time is costly to the inmate. What economics has to add is that, when it comes to drug dealers, some of those costs are passed on to their customers. In other words, one of the likely consequences of reducing prosecution and sentencing of

[48] Chicago Tribune Editors (2006).
[49] In 2008 DEA declared an important fentanyl ingredient as a List I chemical, subjecting it to additional regulation. In 2010, it would prohibit a more finished version of the chemical. Both of these rules are cited in Table 1.
[50] U.S. Department of Justice (2013, 2014) and Federal Bureau of Prisons (2020). See also U.S. Department of Justice, Office of the Inspector General (2017) and Higham, Horwitz and Zezima (2019).

nonviolent drug dealers is that drug dealers would shift to less violent methods in order to be protected by the Holder memo. An example is the shift toward fentanyl mailed from China to local dealers and away from heroin grown in Mexico and traditionally distributed by violent Mexican crime cartels (Pardo, 2019, p. p. 69). As dealers achieve lower costs by doing so, additional likely consequences are that the products they sell would become cheaper, or higher quality, and consumers would respond by purchasing a greater morphine-equivalent quantity (Becker, Murphy, & Grossman, 2006). Additional price reductions may also have come from a shift of competition among sellers away from violence to the price competition that more closely resembles licit markets.

More systematically, the Federal Bureau of Prisons data show that the number of inmates peaked at 219,298 in 2013. Within three years, the number had fallen 27,128 even though it had increased in each of the 33 prior years. A longer version of this paper estimates that this change in prison population reduced costs in illicit markets, of which $2.0 billion annual were savings in the markets for Im opioids and illicit/diverted Rx opioids (Mulligan, 2020). By this estimate, federal sentencing reform may be the single largest subsidy (tax reduction) going toward opioid transactions involving non-medical use among those listed in Table 1 and by itself to reduce Im opioid prices almost 20 percent.[51]

To be clear, the cost and expenditure data from illicit markets are imperfect. Even taken at face value, they suggest that law enforcement by itself may not be enough to reduce Im prices between 2013 and 2016 by the 31 percent indicated in Table 2. Technological change, such as new supply chains, has also been a factor reducing quality-adjusted heroin prices (Farrell, Mansur, & Tullis, 1996; Reuter & Kleiman, Risks and prices: An economic analysis of drug enforcement, 1986; Office of National Drug Control Policy, 2020). Technology is thought to have especially progressed in illicit fentanyl markets. The progress included a series of on-line publication of more efficient synthesis methods, particularly between 2005 and 2013, that "render[ed] a much broader set of individuals qualified and capable of producing fentanyl."[52]

[51] In the tradition of Friedman (1966), this is a positive rather than normative paper. Even a causal interpretation of the correlation between federal sentencing policy and the price and quantity of illicitly-manufactured opioids is consistent with opposite normative conclusions.
[52] Pardo (2019, p. 64). The timing is not necessarily explained by internet use, which had also been cited as a factor in earlier fentanyl episodes (Hempstead and Yildirim 2014).

Twenty-three new fentanyl analogs were reported by U.S. law enforcement between 2016 and 2018.[53]

It is possible that an increase in demand for Rx opioids contributed to the later prevalence of fentanyl in illicit markets, as Powell and Pacula (2021) conclude on the basis of interstate comparisons. Some consumers develop a tolerance for opioids and might welcome (even at the same price) a more potent product like fentanyl, which is often preferred even in legitimate medical practice for treating pain in opioid-tolerant patients (Stanley, 2014). However, the facts that heroin use fell after 2015 or 2016 (Figure 3) and has reached essentially zero in some regions, even while many users prefer heroin to fentanyl at the same price (Pardo, 2019, p. p. 101), suggests that fentanyl supply is also an important factor.

In May 2017, the new Attorney General Jeff Sessions reversed Holder's sentencing policy (Sessions, 2017) but not the downward trend for the prison population (Federal Bureau of Prisons, 2020). Another round of DEA rules intended to curtail fentanyl production and distribution began in May 2016. Ten DEA final rules cited in Table 1 prohibited various fentanyl analogs by putting them on Schedule I.[54] China, which had been a principal source country, began class-wide control of fentanyl analogs in May 2019.

The quarterly opioid overdose fatality rate series also shows a reversal in 2017 Q2. The rate fell in each of the three quarters 2017Q2 through 2017Q4 after increasing in 10 of the previous 11 quarters. The rate fell again in two of the four quarters of 2018. A reversal is also obvious in the Im fatality series, which fell in 2017Q2 after increasing in 21 of the previous 22 quarters. The Im fatality rate was essentially constant in 2017Q3 and fell again in two of the next five quarters. As shown in Figure 3, recent heroin usage as measured by NSDUH shows some of the largest year-over-year drops in the year following the reversal of the Holder memo. However, Figure 4 shows that fatalities from synthetic opioids continued to increase.

[53] In 2017 and 2018, U.S. law enforcement and the Department of Treasury began to prosecute and sanction Chinese nationals found trafficking fentanyl in the U.S. (O'Connor 2018, p. 5).
[54] Those rules are 81 FR 29492, 81 FR 85873, 82 FR 20544, 82 FR 26349, 82 FR 32453, 82 FR 49504, 83 FR 469, 83 FR 4580, 83 FR 5188, and 83 FR 61320. In addition, thiafentanyl, which is used for animals and is roughly 1,000 times more potent than heroin, was put on Schedule II in 2016 (81 FR 58834). Three more rules put three less potent opioid substances on Schedule I (81 FR 22023, 81 FR 79389, and 82 FR 32453).

VII. Quantitative Predictions from the Sufficient Statistics Approach

Although the effects of an opioid policy may change magnitude and direction depending on the status of substitutes and complements, the sufficient statistics approach helps apply specific empirical findings to other contexts. Take the findings cited in Section V.B. that the reformulation of Oxycontin increased Im fatalities in the years 2011-13 almost exactly the same amount it decreased Rx fatalities. This suggest that, on average 2011-13 among those who would have been consuming OxyContin, equation (3) evaluated to essentially zero. By comparing those years to other years or various subpopulations on the basis of r, CROSS, and the relative price (and to some extent ARC and POINT although they tend to cancel in equation (3)), predictions can be made as to whether the sign is positive or negative. With data on r and the relative price, the OxyContin findings can also be taken as providing an estimate of CROSS (the elasticity of Im consumption with respect to the Rx price) that can be applied in other contexts were r and the relative price take on other values. The estimate of CROSS depends on what the Oxycontin consumers on the margin of Im were paying per MGE, which I treat as a parameter for the purposes of assembling Table 2. At a lower price of \$273 per MGE, the implied CROSS is 2.3.[55] At a higher price, corresponding with the price that would be paid on the secondary market, the implied CROSS is 1.4.[56] Intermediate values for CROSS would be found with intermediate values of the OxyContin price point.

The remaining columns of Table 2 show elasticities with respect to a Rx price change that is uniform for all Rx products. Each elasticity is calculated using the sufficient-statistics version of equations (2) or (4) in elasticity form, with measures of r and p_R/p_I taken from the top of the Table, CROSS taken from the left column, and the POINT elasticity assumed to be constant along the entire arc and equal to -0.5.[57] One important result in panel A, especially

[55] The \$273 is the same proportion of the FDA Rx price series for 2011-13 as the \$336 I found for the year 2014, as discussed in Section IV.A. of this paper.

[56] To assess orders of magnitude of CROSS, note that Stehr's (2005) estimates of illicit cigarettes in the U.S. finds, in my notation, that CROSS*$(1-r)/r$ is about 0.7 (as compared to 0.8 to 1.3 in Table 2) even though (i) the illicit sector's share is much less in cigarette markets than in opioid markets and (ii) Stehr did not focus on heavy smokers.

[57] As a close approximation, alternative values of POINT rescale all of the entries proportionally. The panel A entries are calculated with the formula (3) multiplied by $-$POINT r in order to provide a magnitude rather than just a sign. The panel C entries are calculated as $(1 -$

noticeable in the columns for years 2011-13, is that raising the full price of higher-priced OxyContin has a different effect than uniformly raising the full price of all Rx opioids. The Oxycontin effect is assumed to be zero on average in 2011-13 whereas the elasticity with respect to a uniform increase is about −0.35. Although the precise magnitudes are necessarily uncertain due to uncertainty about how the OxyContin price point fits in the overall structure of opioid prices paid by consumers on the Rx-Im margin, the robust lesson is that policies that target higher Rx prices have meaningfully different effects than policies applicable to all Rx-opioid purchases. The difference is in the direction of a larger "jump" term for the targeted policies.

Another lesson from panel A is that the sign of the effect of Rx prices on overall opioid consumption changes over time as Im opioid prices change relative to Rx prices. The sign switches from negative to positive in 2015 for both of the values of CROSS used in the table; an Rx policy that once reduced overall opioid consumption more recently would increase it.[58] Panel B shows the estimates for the Rx-price elasticity of overall consumption "among the treated" (Heckman & Vytlacil, 2001), $\frac{\partial \ln D(p_R,p_I)}{\partial \ln p_R}\frac{1}{r}$. Any gap between the entries in Panel B and POINT (−0.5) reflect the contribution of the jump term, which is negative in years when Rx is cheaper than Im and positive when Rx is more expensive. Panel C shows the effects of Im prices, which for the years shown are always in the direction of less overall consumption.

Arguably the Rx share of opioid-involved deaths is less than the Rx share of opioid consumption among opioid misusers. Panel D pertains to this case, assuming that Im consumption is 33 percent more lethal per MGE than Rx, and showing the resulting effect on overall opioid deaths rather than opioid consumption. The lessons from panel A largely still apply, although the switch from negative to positive sign may occur a little earlier in panel D. On one hand, a lesser value of CROSS is needed to fit the reformulation studies because fewer switchers can nonetheless disproportionately generate fatalities. On the other hand, each switcher from Rx to Im has a larger effect on aggregate mortality in panel D's model than in

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$r)POINT\left[1+\frac{ARC}{POINT}CROSS\,\frac{H(p_I)}{H(p_R)}\left(1-\frac{p_I}{p_R}\right)\right]$, which is derived from equation (4) in the same way.

[58] The sign change would be a year later if CROSS were 1.0, or two years later with CROSS = 0.5.

panel A's. Panel E shows that the sign change occurs a year later if Im prices are 50 percent greater than shown by the estimates at the top of the table.

VIII. Conclusions

Although federal policies toward opioid markets have changed in dozens of ways, their effects can be understood on the basis of just a few characteristics. The first is whether a policy primarily affects prescription opioids (Rx monitoring programs are such a policy), illicitly-manufactured opioids (law enforcement), or both (subsidies for benzos). The second characteristic is the pricing context in which the policy applies. When heroin and fentanyl prices are low enough, as they appear to have been since 2014, Rx policies have the opposite effect on total opioid consumption than they did in an earlier era when Im opioids were more expensive than Rx opioids.

Pricing context within a segment also matters. For example, "prescription takeback" programs affect the secondary market where prescription (Rx) opioids are most expensive. The 2010 Oxycontin reformulation affected brand-name primary-market Rx opioids, which are more expensive than the average Rx opioid purchased at pharmacies, whereas monitoring programs and uniform excise taxes affect all Rx. Rescaled by the amount that they elevate the average full price of Rx opioids, the takeback, reformulation, and monitoring programs are therefore expected to have progressively more negative (or less positive) effects on overall opioid consumption because the average drop in their marginal prices resulting from a switch to illicitly-manufactured (Im) opioids is progressively less for the three groups of affected customers.

This is also the first paper to comprehensively catalog the dozens of federal opioid policy changes, which it identifies in twenty years of the Federal Register and from the Department of Justice press releases. Table 1 helps to begin assessing the third essential policy characteristic, which is the magnitude of the subsidies, taxes, or compliance costs created by policy. An overall pattern that emerges from the table is that opioid policies subsidized and facilitated prescriptions from the year 2000 until about 2010. These policies include "pain standards," DEA rules affecting a health provider's marginal cost of prescribing opioids, and federally-subsidized insurance coverage. Later Rx regulations were tightened while the war on illegal drugs was relaxed. Perhaps not coincidentally, illicitly-manufactured opioids exploded after 2013 as fentanyl came into U.S. markets for the first time on a large scale. The magnitude and characteristics of this explosion are further supported with new quantity metrics provided in this paper. In terms of Figure 1, the broad pattern was a shift from allocation A to B between 2000 and 2013 followed by a shift from B to C in the subsequent years.

A key insight from the economic framework is that there is no such thing as "the effect" of an Im policy or "the effect" of an Rx policy. The direction and magnitude of these effects depend on the price differentials between the Rx and Im segments of the market more than they depend on the price sensitivity of total opioid demand. Prior to 2013, Im opioids were more expensive on the margin, by as much as a factor of five in the 1990s. I estimate that by 2017 Im opioids cost, on average, just a third at retail of what Rx opioids cost due to the presence of illicitly-manufactured fentanyl. Policies that increase the full price of Rx opioids in 2017 are likely increasing overall misuse, even though the same policy might have significantly reduced overall misuse a decade earlier. Nevertheless, the economic framework can be expressed in a sufficient statistics format that facilitates the application, as in Table 2, of specific empirical findings to other contexts in which relative prices are quite different.

Guided by the economic framework, this paper also assembles empirical evidence that sheds light on the roles of habit, addiction, and drug tolerance. That prescriptions can be a gateway to consumption of heroin or fentanyl, is supported by the life cycle of opioid misuse by type of opioid, especially during the formative ages 14-25. Adolescents (aged 14-17), who tend to be comparatively Rx-intensive, are the only group where overdose deaths involving Rx opioids fell enough after 2010 to more than offset increases in deaths involving heroin and synthetic opioids. On the other hand, comparisons of whites and blacks suggest that a history of Rx-opioid misuse need not precede an increase in fatalities from synthetic opioids. Although blacks, especially young black men, have been less Rx-intensive in their opioid misuse, by 2019 their overall fatality rate from opioids exceeded that of whites.

The same approach has potential for resolving debates in the literature as to the mortality effects of policies that increase the supply of naloxone, which is "a medication designed to rapidly reverse opioid overdose" (National Institute on Drug Abuse, 2018). There is not only the question of whether a life-saving effect of naloxone holding opioid abuse constant is offset by an increase in the amount of abuse (Doleac & Mukherjee, 2018; Rees, Sabia, Argys, Dave, & Latshaw, 2019), but also the composition of that abuse between Rx and Im. Even if naloxone had the same mortality effect per equivalent quantity of Rx as Im, the fact that it increases the quantity of opioid consumption by itself encourages a shift from Rx to Im where marginal prices are now lower.

Dowell et al (2017)'s analysis of life tables shows that overdose deaths involving opioids subtracted 0.21 years from aggregate life expectancy between 2000 and 2015, which is a period when the annual opioid fatality rate increased by 7.7 per one hundred thousand population (see also Currie and Schwandt 2020). Although their calculations have not yet been updated for the entire period 2013-2017 when Im opioids became so cheap, we know that the annual rate increased by 6.8 per one hundred thousand and can therefore estimate that overdose deaths also subtracted about 0.2 years from life expectancy during that time frame. By comparison, the actual change in life expectancy was –0.20. In other words, life expectancy would have fallen far less, if at all, if the opioid fatality rate had remained at 2013 levels, which were already elevated by historical standards.

There is much more to 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. Better predictions would be possible with estimates of short and long run supply elasticities, and how they are different for heroin and fentanyl. Something akin to technological change or expanded international trade has at times been an important driver of illegal drug prices. For example, real heroin prices fell by a factor of three in the 1990s (United Nations Office on Drugs and Crime, 2018). As long as costs remain low, nonmedical use of opioids and other illegal drugs may never return to the lower levels that they once were.

IX. Appendix I: Additional Consumer-Theory Results

IX.A. Aggregate Opioid Demand with Many Dimensions of Heterogeneity

Let the vector $\theta \in \Theta$ index consumer characteristics that affect preferences and potentially add an idiosyncratic component to income In this way, consumers differ, among other things, in terms of the level of and elasticity of demand, perhaps reflecting heterogeneous consumption histories. The consumers on the margin between Rx-only and Im-only are those with the critical fixed cost f^* satisfying $v(p_I, y - f^*(p_R, p_I; \theta); \theta) = v(p_R, y; \theta)$, where v is the same indirect utility function referenced in the main text except now it is indexed by θ. As in the main text, mass points in the joint distribution of f and θ are ruled out; let $g(f,\theta)$ denote the density function.

Define a marginal consumer's *ARC* elasticity as:

$$ARC(p_R, p_I; \theta) \equiv \frac{1 - \dfrac{M(p_R, y; \theta)}{M(p_I, y - f^*(p_R, p_I; \theta); \theta)}}{1 - \dfrac{p_R}{p_I}} < 0$$

where M denotes the Marshallian demand for opioids corresponding to the indirect utility function v. *ARC* is a Hicksian elasticity because of the income compensation. Let $ARC(p_R,p_I)$ denote the average *ARC* among all marginal consumers:

$$ARC(p_R, p_I) \equiv \frac{\int_\Theta ARC(p_R, p_I; \theta) g(f^*(p_R, p_I; \theta), \theta) d\theta}{\int_\Theta g(f^*(p_R, p_I; \theta), \theta) d\theta} < 0$$

Define a consumption-weighted average Marshallian point elasticity among Rx consumers:

$$POINT(p_R, p_I) \equiv p_R \frac{\int_\Theta \int_{f^*(p_R,p_I;\theta)}^{\infty} M_p(p_R, y; \theta) g(f, \theta) df\, d\theta}{\int_\Theta \int_{f^*(p_R,p_I;\theta)}^{\infty} M(p_R, y; \theta) g(f, \theta) df\, d\theta} < 0$$

where M's p subscript indicates the first partial derivative with respect to price. The main-text definitions of ARC and POINT are the special case in which individuals are homogeneous in terms of these elasticities. CROSS is an aggregate of the Im responses to p_R, weighting each marginal consumer by his ARC elasticity and expressed in elasticity format:

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

The generalizations of equation (1) and its price derivative are, respectively:

$$D(p_R, p_I) = \int_\Theta \int_{-\infty}^{f^*(p_R,p_I;\theta)} M(p_I, y-f; \theta) g(f,\theta) df\, d\theta + \int_\Theta \int_{f^*(p_R,p_I;\theta)}^{\infty} M(p_R, y; \theta) g(f,\theta) df\, d\theta \tag{5}$$

$$\frac{\partial D(p_R, p_I)}{\partial p_R} = \int_\Theta [M(p_I, y - f^*(p_R, p_I; \theta); \theta) - M(p_R, y; \theta)] g(f^*(p_R, p_I; \theta), \theta) \frac{\partial f^*(p_R, p_I; \theta)}{\partial p_R} d\theta + \int_\Theta \int_{f^*(p_R,p_I;\theta)}^{\infty} M_p(p_R, y; \theta) g(f,\theta) df\, d\theta \tag{6}$$

As in the main text, let 1–*r* and *r* denote the shares of the first and second terms in (5), respectively. Eliminate the demand derivatives and differences from equation (6) using the definitions of *ARC* and *POINT*:

$$\frac{\partial D(p_R, p_I)}{\partial p_R} = \left(1 - \frac{p_R}{p_I}\right) \int_\Theta ARC(p_R, p_I; \theta) M(p_I, y - f^*(p_R, p_I; \theta); \theta) g(f^*(p_R, p_I; \theta), \theta) \frac{\partial f^*(p_R, p_I; \theta)}{\partial p_R} d\theta + \frac{POINT}{p_R} \int_\Theta \int_{f^*(p_R,p_I;\theta)}^{\infty} M(p_R, y; \theta) g(f,\theta) df\, d\theta$$

Equation (3) is obtained by factoring out the final term, using the definitions of CROSS (after evaluating its partial derivative) and *r* to eliminate the integrals, and suppressing the dependence of elasticities on prices:

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

IX.B. Additional Properties of the Consumer Budget Set

The consumer's budget constraint is piecewise linear in the $[Q,z]$ plane, formed as the upper envelope of the three linear budget constraints corresponding to the three possible decisions regarding fixed costs: $y = z + f_R + Q\ p_R$, $y = z + f_I + Q\ p_I$, and $y = z + f_R + f_I + Q\ E(1,p_R,p_I;0,0)$, respectively. Assuming that $f_I > f_R > 0$, 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 $x = q_R$ and $y = q_I$. The upper envelope consists of only one piece if $p_R/p_I < Q_{10}$ (Figure 1a). It consists of two pieces (as in Figure 1b) if and only if $p_R/p_I \geq Q_{01}$ or $Q_{10} < p_R/p_I \leq 1$. In the former (latter) case, the mixed (Im-only) constraint is dominated by the other two, respectively. The remaining interval is where three pieces are possible, with the mixed piece forming the upper envelope at the highest quantities. When Rx and Im are close substitutes, the gap between Q_{10} and Q_{01} is less and the likely cases are either two pieces or one.

IX.C. The Role of Income Effects

The applicability of rational choice to the demand for addictive drugs is a matter of vigorous debate. The argument against notes that drugs cause "persistent changes in the brain structures and functions known to be involved in the motivation of behavior" and that frequently "the addict expresses a desire not to consume drugs prior to, after, or even during the drug intake" (Henden, Melberg, & Rogeberg, 2013). However, both sides of the debate acknowledge budget constraints, which are at least half of the rational choice model of consumer behavior. Using a random demand model with linear budget constraints, Becker (1962) shows that market demand for a commodity slopes down even when the market lacks any rational consumer.[59] For example, an addict spending all (or any fixed share less than one) of his income on a drug would have unit-elastic demand: doubling the price would require him to reduce consumption by a factor of two.

[59] Conversely, the fact that an addict reduces drug consumption in response to higher prices is not proof of rationality (Becker and Murphy 1988).

Figures 10 and 11 extend Becker's constant-budget-share analysis to the two-part budget constraint. As the Rx opioid price varies by itself, two levels of consumption of all other goods are consistent with a single budget share, depending on whether opioids are sourced from Rx or Im.[60] These two levels are shown as horizontal blue lines in Figure 10. I assume that consumers do not choose a dominated point on the budget constraint (that is, a point where more of both goods can be purchased), and achieve the fixed budget share in the Rx market whenever the dominance criteria does not distinguish between Rx and Im. Beginning with a low Rx price (less than p^*), marginal increases in price reduce opioid consumption with an elasticity of negative one, moving the allocation horizontally along Figure 10's upper horizontal line toward allocation B. As the price passes p^*, opioid consumption jumps up from q_B to q_C. Further increases in the Rx price have no effect on opioid consumption because opioids are sourced from Im.

Figure 11 shows this "irrational" demand curve in quantity-price space together with the rational demand curve for a consumer having Cobb-Douglas preferences with share parameter equal to the irrational consumer's fixed budget share. For Rx prices less than $p_L < p^*$, both theories of demand generate the same a unit-elastic curve. For Rx prices greater than p^*, both generate the same fixed quantity q_C. The difference is that the rational consumer switches from Rx to Im before Rx is strictly dominated. The irrational consumer holds out until the Rx price exceeds p^*, at which point the discrete increase in opioid consumption is greater than the increase made by the rational consumer at price p_L.[61] Using only income effects and the idea of dominance, Figures 10 and 11 thereby strengthen the unconventional prediction that increasing Rx opioid prices increases opioid consumption over the range in which consumers switch from Rx to Im. Moreover, the properties of the aggregate demand function (1) used in this paper do not depend whether consumers are assumed to be "rational" as in Figure 1 or "irrational" as in Figure 10.

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[60] Here I assume that the budget share is applied to income net of the black-market fixed cost. The net income would be disposable income if part of the fixed cost accrued in the form of a monetary fine, less employment, or lower-paying work.

[61] In other words, relative to the rational consumer, the jump in the irrational consumer's opioid consumption is disproportionate to the sensitivity of opioid consumption to the Rx price at prices below. The rational consumer's jump from point B to C in Figure 1 is just a movement along a single Hicksian demand curve, whose substitution properties are reflected in the Marshallian demand curve shown in Figure 9 for Rx prices below p_L.

IX.D. Adding Safety Differentials to the Demand System

The main text of the paper uses equation (1) to analyze fatal rates under the assumption that nonmedical opioid consumption, measured in morphine-gram equivalents (MGEs) is proportional to fatality rates. The purpose of this appendix is to extend the model to include the common assertion (Drug Enforcement Administration, 2016; Frank & Pollack, 2017; Ciccarone, Ondocsin, & Mars, 2017) that Im opioid consumption is more dangerous per MGE than Rx opioid consumption.

Let $m \geq 1$ denote the extra mortality associated with each Im MGE consumed and $\mu \leq 0$ denote the elasticity of MGE demand with respect to m, holding constant the retail price. The modified model of opioid consumption becomes:

$$D(p_R, p_I) = F(p_R, p_I, m)m^{\mu}H(p_I) + [1 - F(p_R, p_I, m)]H(p_R) \tag{7}$$

The corresponding model of opioid fatalities is, up to a factor of proportionality:

$$\widehat{D}(p_R, p_I) = F(p_R, p_I, m)m^{1+\mu}H(p_I) + [1 - F(p_R, p_I, m)]H(p_R) \tag{8}$$

Note that the safety differential has offsetting effects on Im mortality, so that the mortality elasticity with respect to m, can have either sign. Indeed, as revealed by policies to enhance the safety of opioid abuse, $1+\mu$ is close enough to zero that researchers cannot agree on its sign (Doleac & Mukherjee, 2018; Rees, Sabia, Argys, Dave, & Latshaw, 2019). If $1+\mu$ were exactly zero, then the model (8) is essentially the same as the model (1) used in the main text, except that equation (1) would have to be interpreted as calculating "mortality-gram equivalents" rather than morphine-gram equivalents.[62]

The model (8) has three terms in its formula for the elasticity of aggregate mortality with respect to the Rx price:

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[62] Table 1 assumes that changes over time in F are due to price changes rather than nonpecuniary factors, which is consistent with the model (8) if m is constant over time. To the extent that the safety differential increased over time, Table 1 understates [check] the effects of prices on F because they are partially offset by changes in m.

$$\frac{\partial \ln \widehat{D}(p_R, p_I)}{\partial \ln p_R} = -r\,\eta - (1-r)ARC\frac{CROSS}{m^{1+\mu}}\left(\frac{p_R}{p_I} - 1\right) + (1-r)CROSS\frac{m^{1+\mu} - 1}{m^{1+\mu}} \tag{9}$$

where *r* now denotes the share of opioid fatalities from Rx. As before $\eta < 0$ denotes elasticity of source-conditional demand *H*(), ARC < 0 denotes the arc elasticity of *H* with respect to the two prices, and CROSS > 0 the elasticity of *F* with respect to the Rx price.[63] The final two terms are both positive if $p_R > p_I$ and $m^{1+\mu} > 1$ and either can exceed the first term in magnitude. In particular, even with no price differential, Rx prices can increase total mortality even though it would not increase total opioid consumption. The elasticity formula reduces to the formula used in the main text when $m^{1+\mu} = 1$.

Note that the additional danger of Im opioid consumption ($m > 1$), relative to Rx opioid consumption, is not apparent in the aggregate data. Take the year 2010, prior to the prevalence of fentanyl, in which 13,903 persons died from overdose involving Rx opioids. Because total Rx opioid sales were 247 million MGEs, and many of those were not abused (or even consumed), there were at least 56 fatal Rx overdoses per million Rx MGEs consumed, and probably closer to 100 fatalities per million MGEs. During the same year, 2888 overdose deaths involved heroin whereas ONDCP estimates that heroin consumption was 135 million MGEs, or about 21 fatalities per million MGEs consumed. As discussed in connection with Table 4, I believe that ONDCP's demand model exaggerates heroin consumption, but even my rescaled-NSDUH method puts heroin fatalities at 87 per million MGEs consumed in 2010.

Between 2010 and 2016, fatalities involving Im opioids increased sharply as fentanyl came into the market, but so did Im opioid consumption. ONDCP estimates that heroin consumption increased 74 percent during that time. In addition, the heroin was increasingly sold with fentanyl, which has a much greater MGE content. By CEA's method of assessing shares of heroin versus fentanyl, the increase in Im MGEs consumed was 206 percent between 2010 and 2016. The increase in fatalities involving Im opioids was 390 percent during that period, which is consistent with an increase in deaths per MGE by a factor of 1.6.

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[63] The definition of ARC is $\frac{1-H(p_R)/H(p_I)}{1-p_R/p_I} < 0$.

X. Appendix II: Identifying Federal Opioid Policies

I identified federal opioid policies for potential analysis by searching the Federal Register for the years 2001-19 for final rules from the Department of Health and Human Services (HHS), the Department of Justice (DOJ), the Department of Labor (DOL), the Department of Transportation (DOT), the Department of Treasury, and the Department of Veterans Affairs containing the word "opioids."[64] Because the DOJ often uses discretion (i.e., case-specific facts and circumstances) rather than the rulemaking process, I also searched DOJ press releases describing department initiatives with the word "drug" and relating to prosecution and sentencing. In reviewing the results, I followed and reviewed sources cited as requiring rulemaking, the most significant of which were the 2010 Affordable Care Act and the Medicare Modernization Act of 2003. I also reviewed relevant prior rulemaking cited by the rules identified in the search.

A number of the DOJ press releases refer to sentencing of, or judgments against, specific companies, gang members, and international drug smugglers. Since 2017, some of the press releases also refer to new procedures for immigration enforcement. The aggregate of these may be significant, but I left that topic for future research except to the extent that it is reflected in the federal prison population.

I judged a number of final rules to be too insignificant for aggregate analysis, but list them here for completeness. Treasury issued rules, such as 77 FR 64663 (2012), that exclude opioids from the medicine exemption from the Iranian sanctions. DOL and DOT issued rules regarding the possession of opioids on the job for specific occupations. HHS issued rules in 2017, 2018, and 2019, respectively, requiring confidentiality of substance abuse disorder patient records (82 FR 6052), requiring Accountable Care Organizations to monitor opioid utilization (83 FR 67816), and providing guidelines for federal workplace drug testing programs (84 FR 57554). A 2002 rule from the Drug Enforcement Administration (DEA, which is part of the DOJ) moved the semisynthetic opioid buprenorphine from Schedule V to Schedule III, which is

[64] A rule was excluded if "opioids" were mentioned only as part of a summary of public comments rather than the agency's own description of its rule.

more restrictive (67 FR 62354). A 2012 DEA rule (79 FR 37623) placed the opioid Tramadol onto Schedule IV. Tramadol is one-tenth as potent that morphine and one-fifteenth oxycodone.

XI. Appendix III: Time Series for the Relative Price of Rx and Im Opioids

Using IQVIA National Sales Perspectives, FDA (2018) compiled prescription opioid sales and invoice price time series for 1992-2016. The FDA calculations do not reflect third-party payments that put out-of-pocket prices below invoice prices, especially in more recent years. On the other hand, they do not reflect secondary market prices, which are much greater than pharmacy invoice prices (more on this below) but are only available for two years. The FDA calculations are used for my Figure 2, except that generic and branded are reweighted to reflect in 2011 the relative changes in the two types of Rx opioids between 2010 and 2015 as the market was substituting toward Im. The generic weight evolves over time in proportion with the market-wide average generic share as reported by FDA (2018).

The Drug Enforcement Administration (DEA) collects heroin prices as a byproduct of its efforts to prosecute offenders and to monitor the drug supply. The estimated prices vary widely according to the size of the transaction and how (or whether) the sample is reweighted in attempt to represent the average consumer rather than the average prosecution. The reweighted averages tend to be much lower, but are not consistently reported for long periods of time. Moreover, it is known that regular customers receive substantial discounts, which DEA agents are often not receiving (Arkes, Pacula, Paddock, Caulkins, & Reuter, 2008; Jacques, Allen, & Wright, 2014). Another concern for comparing quality-adjusted marginal Rx and Im prices is that Im opioids are not only more potent due to chemical makeup but also that they tend to be administered differently (intravenously).[65] Of course, becoming a regular customer or learning a new method of administration have their costs but for the purposes of my analysis fixed costs of this type must be distinguished from marginal costs.

In order to minimize exaggeration of the marginal heroin price for regular customers, I use the long time series assembled by United Nations Office on Drugs and Crime (2018) from

[65] National Academies of Sciences, Medicine and others (2017, Chapter 4).

DEA and other sources because its heroin prices somewhat less than the long time series published by ONDCP (2019).[66] It is adjusted for fentanyl content using the method of Council of Economic Advisers (April 2019), except that I updated some of its data sources and assumed a more conservative 15-to-1 ratio for the potency of fentanyl versus pure heroin. The adjustment is minimal for most prior years to 2014.[67] By 2018, it reduces the Im opioid price by 66 percent relative to pure and unmixed heroin.

Because the UN heroin prices are still high compared to the reweighted (but not durably available) series from DEA, I suspect that the relative price of heroin shown in Figure 3 and Table 2 is still somewhat exaggerated. The important and more robust conclusion is that heroin was significantly more expensive than Rx in the 1990s and is now significantly cheaper.

[66] The percentage changes over time are similar among the various sources. The exception is after 2015, where DEA has more of an increase in heroin prices per pure gram than other sources do. In order to be conservative as to the reduction in heroin prices in recent years, I splice the DEA series onto the UN series after 2015.

[67] During the aforementioned fentanyl episode of 2006, the adjustment reduces the Im opioid price by 21 percent relative to pure and unmixed heroin.

Table 1. Changes in Federal incentives related to the market for opioids

	Incentives for:		
Year	Prescribers	Patient Rx purchases	Illicit Manafucture and Distribution
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]
2011		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]	*
2016	CARA Act	CARA Act	*
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	*; SUPPORT Act
2019	CMS removes pain management questions from HCAHPS[u]		*; Series of new DEA prohibitions.[v]

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 Existing Requirements Under the Controlled Substances Act for Prescribing Schedule II Controlled Substances. August 2005.
[d]70 FR 4228 (January 2005).
[e]DEA prohibits chemicals used to manufacturer fentanyl 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
[l]Good (2018).
[m]77 FR 22076 (April 2012).
[n]Holder, Eric. "Department Policy on Charging Mandatory Minimum Sentences and Recidivist Enhancements in Certain Drug Cases."
[o]DEA. Schedules of Controlled Substances: Rescheduling Hydrocodone Combination Products from Schedule III to Schedule II. August 2014.
[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-ana
[s]Sessions, Jeff. "Department Charging and Sentencing Policy."
[t]83 FR 32784 (July 2012).
[u]CMS. "HCAHPS Update Training" (February 2019).
[v]Spanning 5/2016 through 11/2019, 10 DEA rules put various fentanyl analogs on Schedule I.
*see also competition from Rx

Table 2. Generalizing the Reformulation Effect to Other Opioid Policies in Other Years

Data			A policy that uniformly increases Rx or Im prices in year: 2011	2012	2013	2014	2015	2016	2017
Marginal Rx price, $2016 per MGE			84.1	87.1	101.0	111.7	130.4	144.1	*144.1*
Marginal Im price, $2016 per MGE			108.7	95.1	96.5	95.2	68.5	66.3	47.4
Rx share of opioid-involved deaths			0.714	0.648	0.569	0.495	0.418	0.331	0.266
Elasticity	**OxyContin price point, 2011-13, $2016 per MGE**	**CROSS**	**2011**	**2012**	**2013**	**2014**	**2015**	**2016**	**2017**
A. Aggregate opioid	273	2.3	-0.45	-0.36	-0.26	-0.16	0.16	0.33	0.58
consumption wrt Rx price	715	1.4	-0.41	-0.35	-0.27	-0.19	0.02	0.15	0.32
B. Treated opioid	273	2.3	-0.63	-0.56	-0.46	-0.32	0.38	0.99	2.20
consumption wrt Rx price	715	1.4	-0.58	-0.54	-0.47	-0.39	0.05	0.44	1.20
C. Aggregate opioid	273	2.3	-0.04	-0.14	-0.24	-0.33	-0.56	-0.67	-0.78
consumption wrt Im price	715	1.4	-0.08	-0.15	-0.23	-0.30	-0.46	-0.55	-0.63
Sensitivity Analysis: Deaths weight Im consumption 33% more									
D. Aggregate opioid deaths	273	1.7	-0.29	-0.20	-0.09	0.01	0.23	0.38	0.56
wrt Rx price	715	1.3	-0.30	-0.23	-0.14	-0.05	0.12	0.25	0.39
Sensitivity Analysis: 50% higher Im prices at all dates									
E. Aggregate opioid	273	2.3	-0.62	-0.55	-0.48	-0.40	-0.06	0.10	0.38
consumption wrt Rx price	715	1.6	-0.54	-0.48	-0.42	-0.35	-0.10	0.01	0.22

Notes: Each table entry is a market-level price elasticity calculated from the data at the top and from CROSS based the paper's model (1)-(4). 2017 Rx prices are assumed to be the same as 2016. POINT is assumed to be -0.5; alternative values would essentially proportionally rescale the table's year-specific elasticity entries. CROSS, which is the only connection between the Table's columns, is derived from the Oxycontin price point column to fit the OxyContin reformulation, with Rx and Im consumption exactly offsetting in aggregate on average 2011-13. Panel B is panel A divided by the Rx share.

Consumption responses to Rx price changes

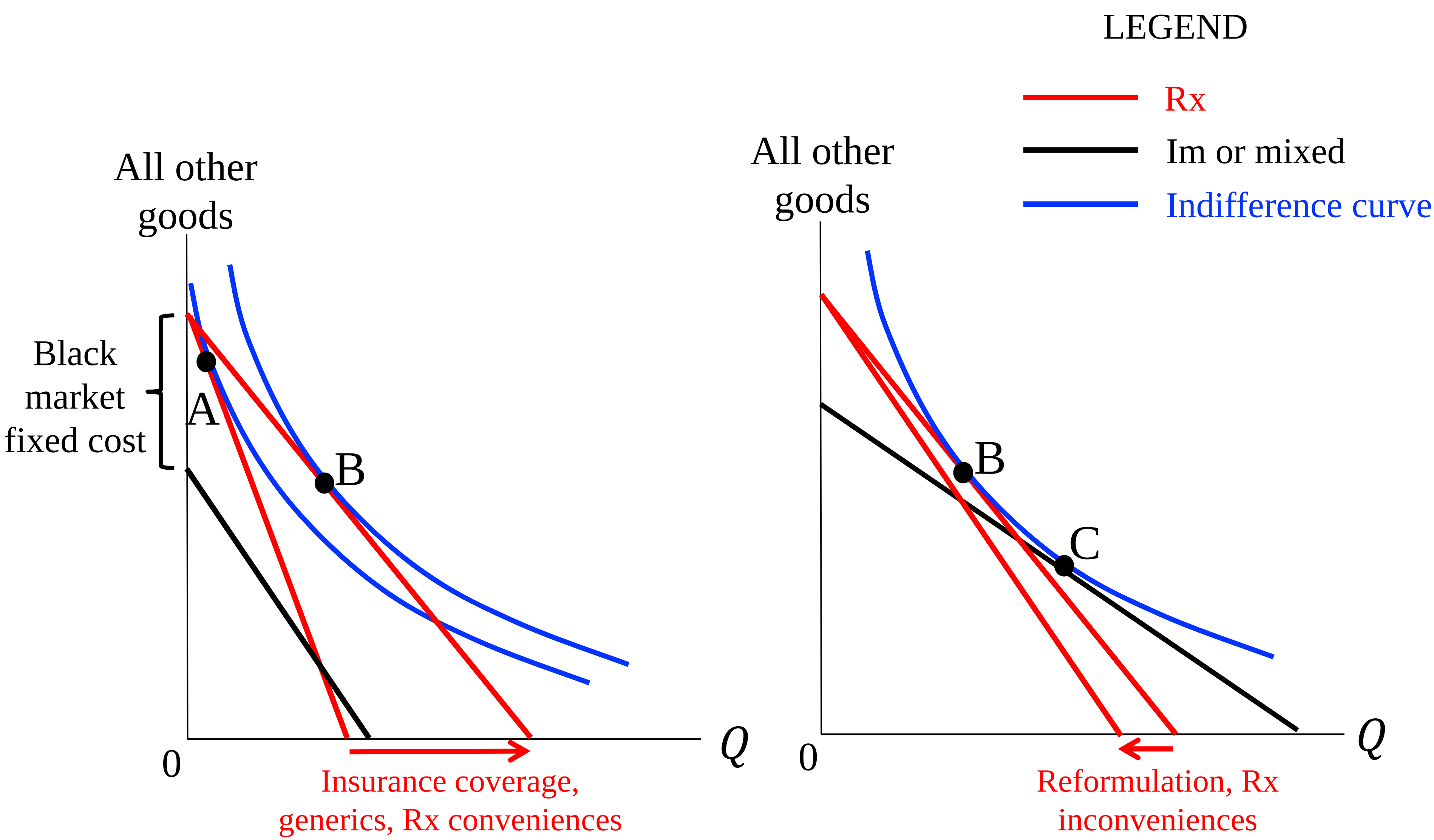
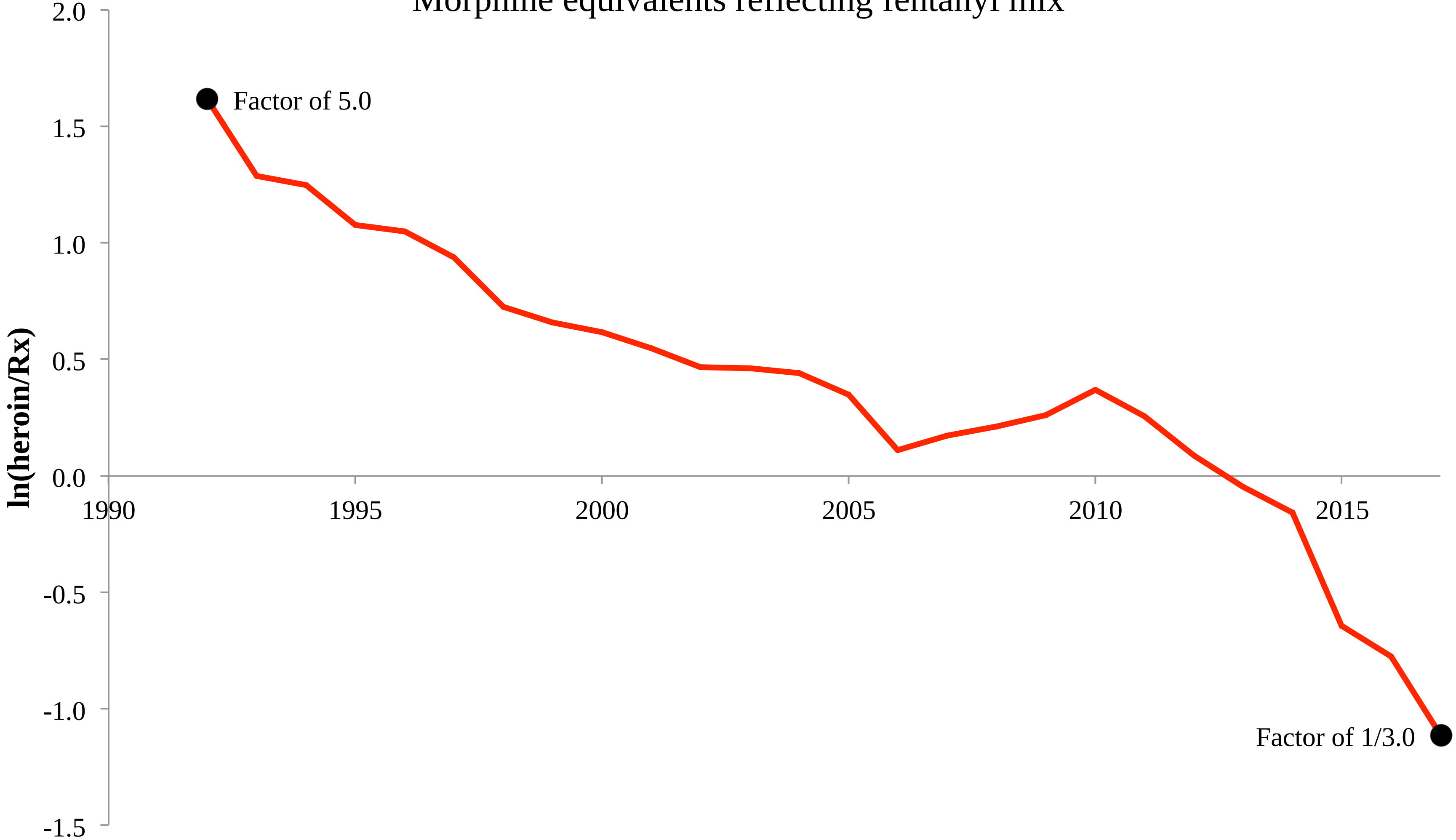


Fig 1a. Im marginal prices are high

Fig 1b. Im marginal prices are low

Figure 2. Retail Prices: Heroin Relative to Rx Opioids

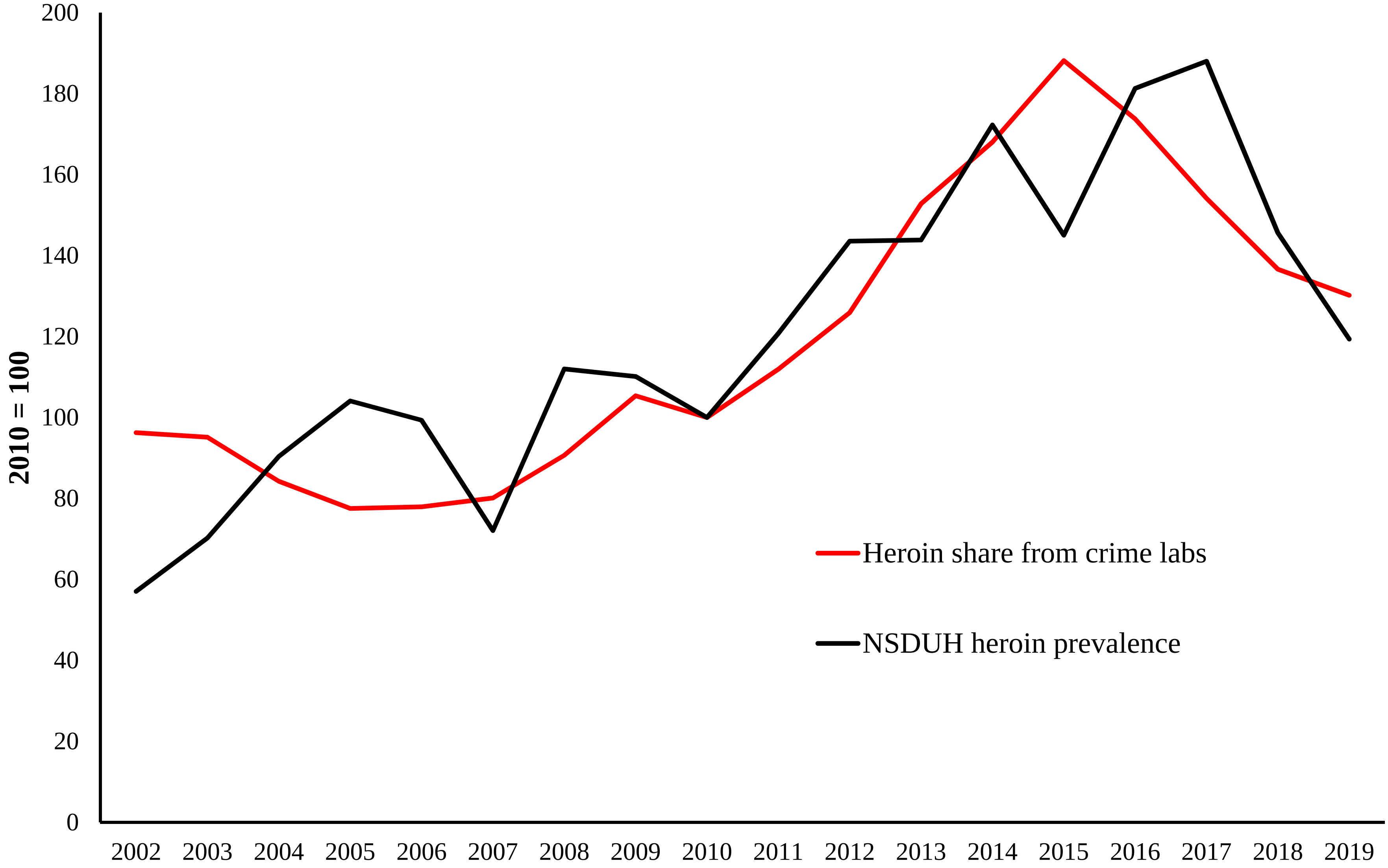
Morphine equivalents reflecting fentanyl mix



Sources: UN, DEA, FDA, CEA.

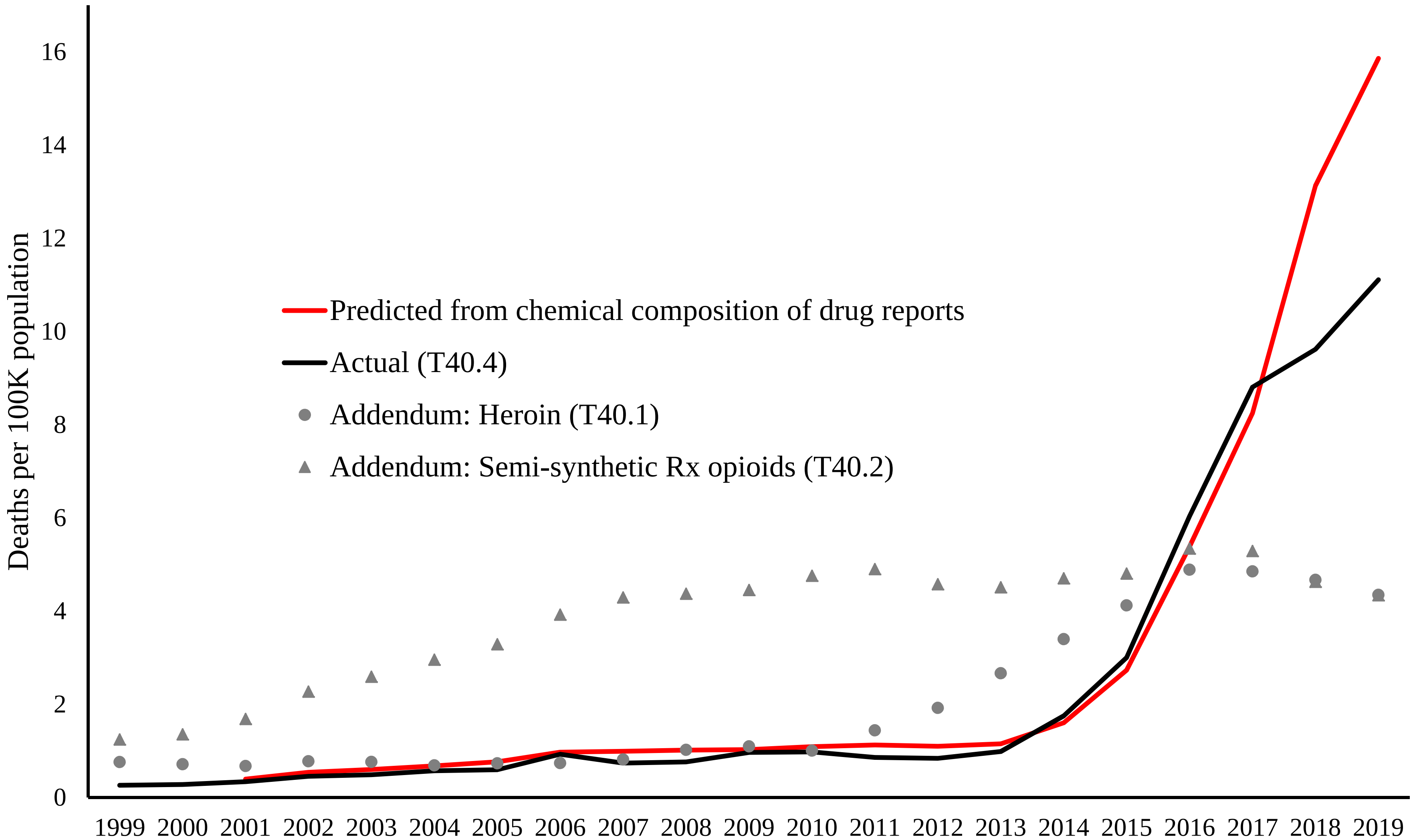
Note: 2011 Rx prices are 29-71 generic vs branded, reflecting their shares in aggregate Rx opioid quantity reductions 2011-15. Rx prices obtained from the secondary market (higher) or net of third-party payment (lower) are not reflected in this series.

Figure 3. NFLIS heroin reports and NSDUH heroin prevalence



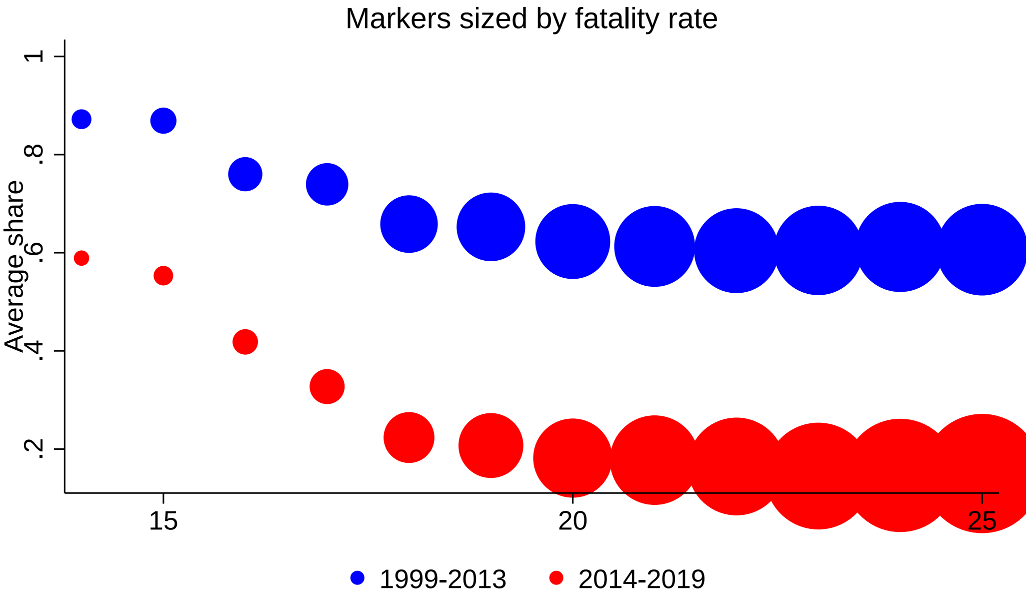
Sources: The crime-lab series is the herion row from the annual NFLIS Drug Reports, Table 1.1. NSDUH series is the variable HERMON, weighted using the 18-year analysis weight.

Figure 4. Rates of Fatal Overdose from Synthetic Opioids



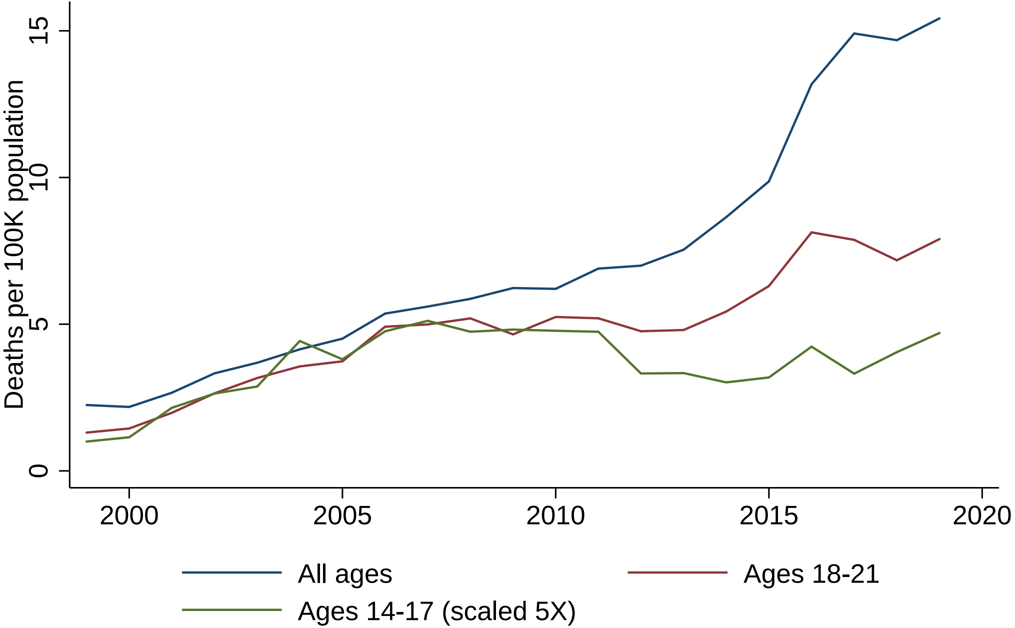
Sources: The T40.1, T40.2, and T40.4 series are from MCOD records with drug poisoning as underlying cause of death. The red series is predicted from heroin deaths (T40.1), semi-synthetic prescription-opioid deaths (T40.2), and the relative frequency of fentanyl and heroin in drug reports from crime labs (NFLIS).

Figure 5. Rx share of opioid fatalities by age and time period



Source: MCOD micro data with drug poisoning as underlying cause of death. Rx shares are conditional on year effects, using 2007 and 2016 as benchmarks. Share numerator is records with multiple-cause T-codes T40.2 and T40.3 (but not T40.0, T40.1, or T40.4) and denominator records with T40.0-T40.4.

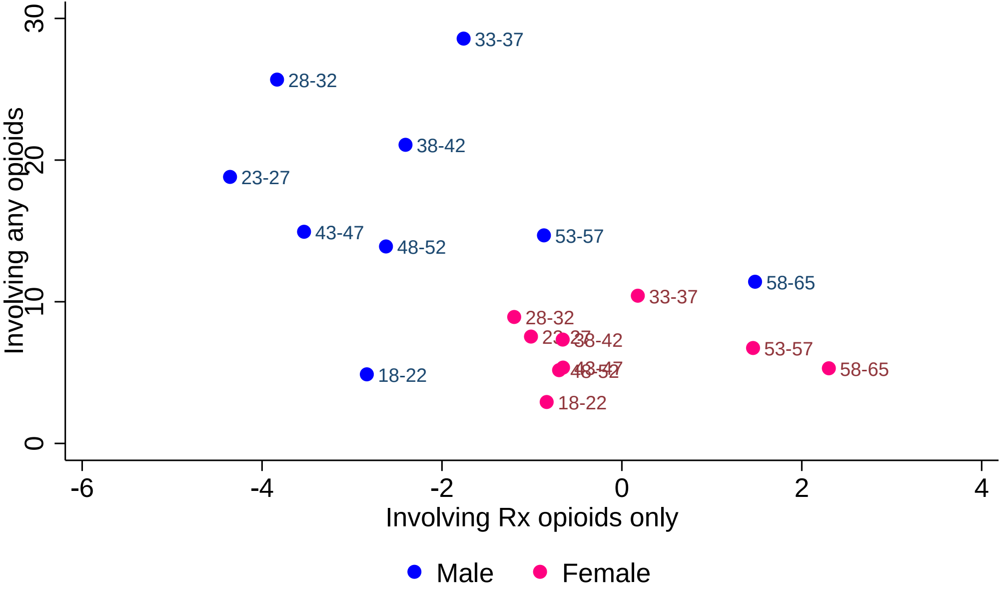
Figure 6. Opioid fatalities by age and year



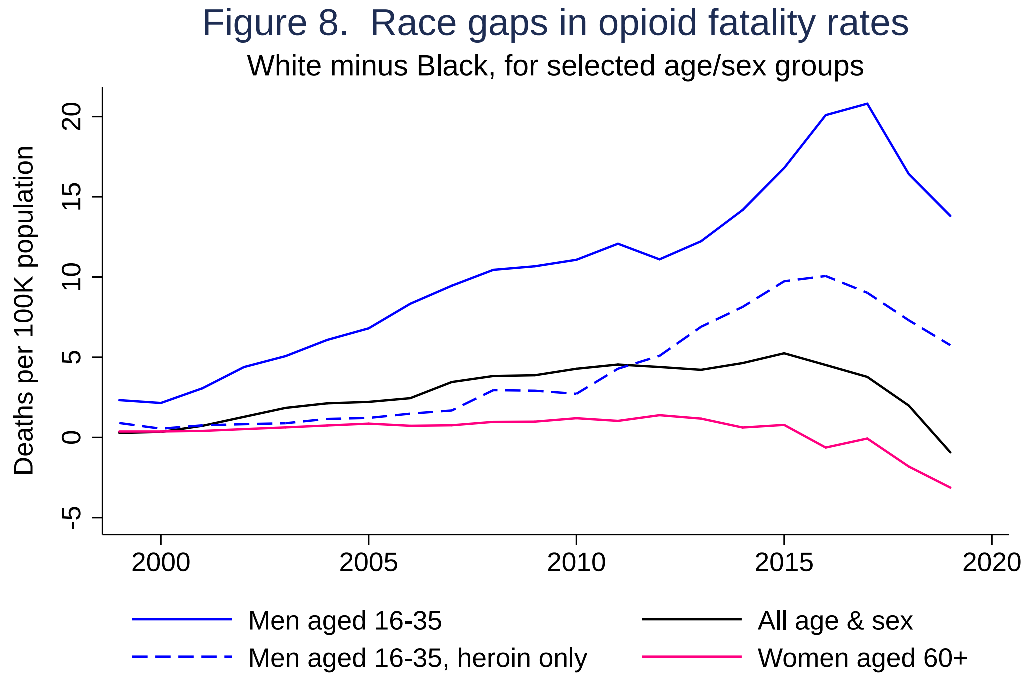
Source: MCOD micro data with drug poisoning as underlying cause of death and multiple causes including any T-code T40.0-T40.4.

Figure 7. Overdose death rate changes 2009-10 to 2016-17

per 100,000 population, ages 18-65

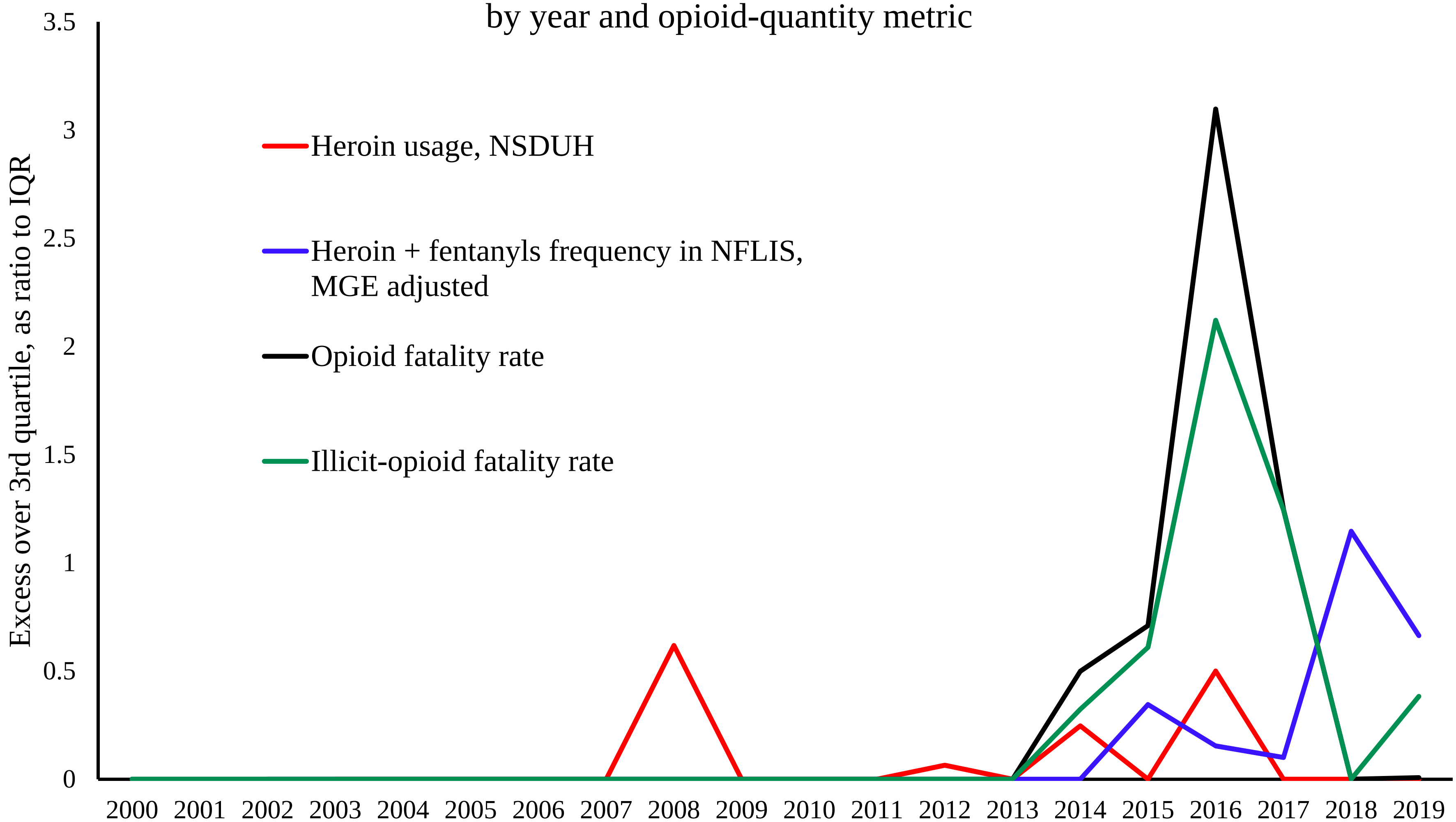


Source: MCOD micro data with drug poisoning as underlying cause of death and multiple causes including any T-code T40.0-T40.4.



Source: MCOD micro data with drug poisoning as underlying cause of death and multiple causes included any T-code T40.0-T40.4.

Figure 9. Right-tail outliers in yearly changes,
by year and opioid-quantity metric



Notes and sources: Zero is a change not in the top quartile of changes. Any change in top quartile is expressed as an excess over the 75th percentile change and then divided by the interquartile range. Fatality rates are from MCOD records with drug poisoning as underlying cause of death and multiple cause with any of the T-codes T40.0-T40.4. NSDUH heroin usage is series HERMON weighted with ANALWC18. NFLIS series are from Drug Reports Table 1.1, with fentanyl weighted by a MGE factor of 3.7 based on customs seizures.

Figure 10. Predicting the "choices" of an irrational consumer

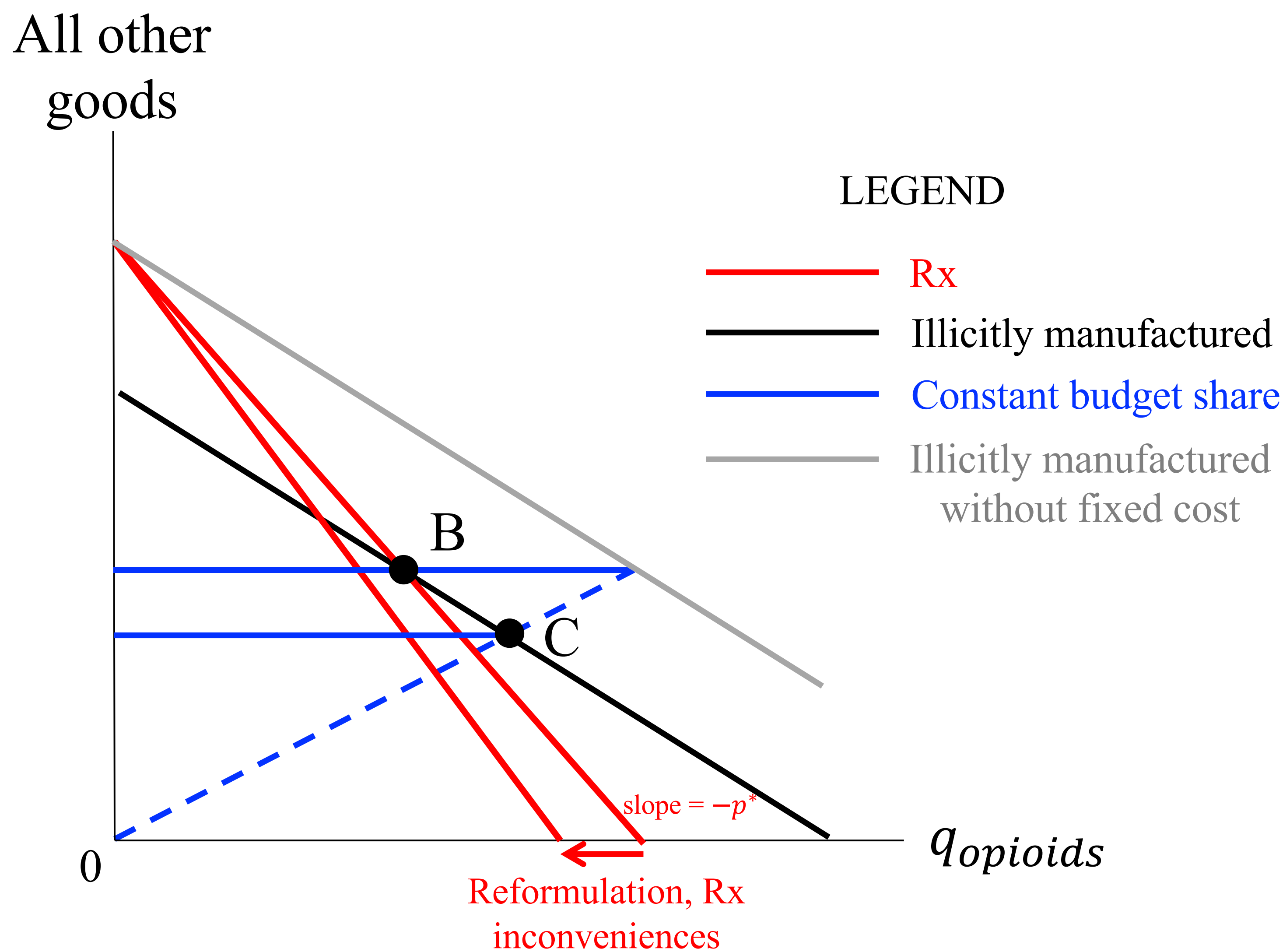
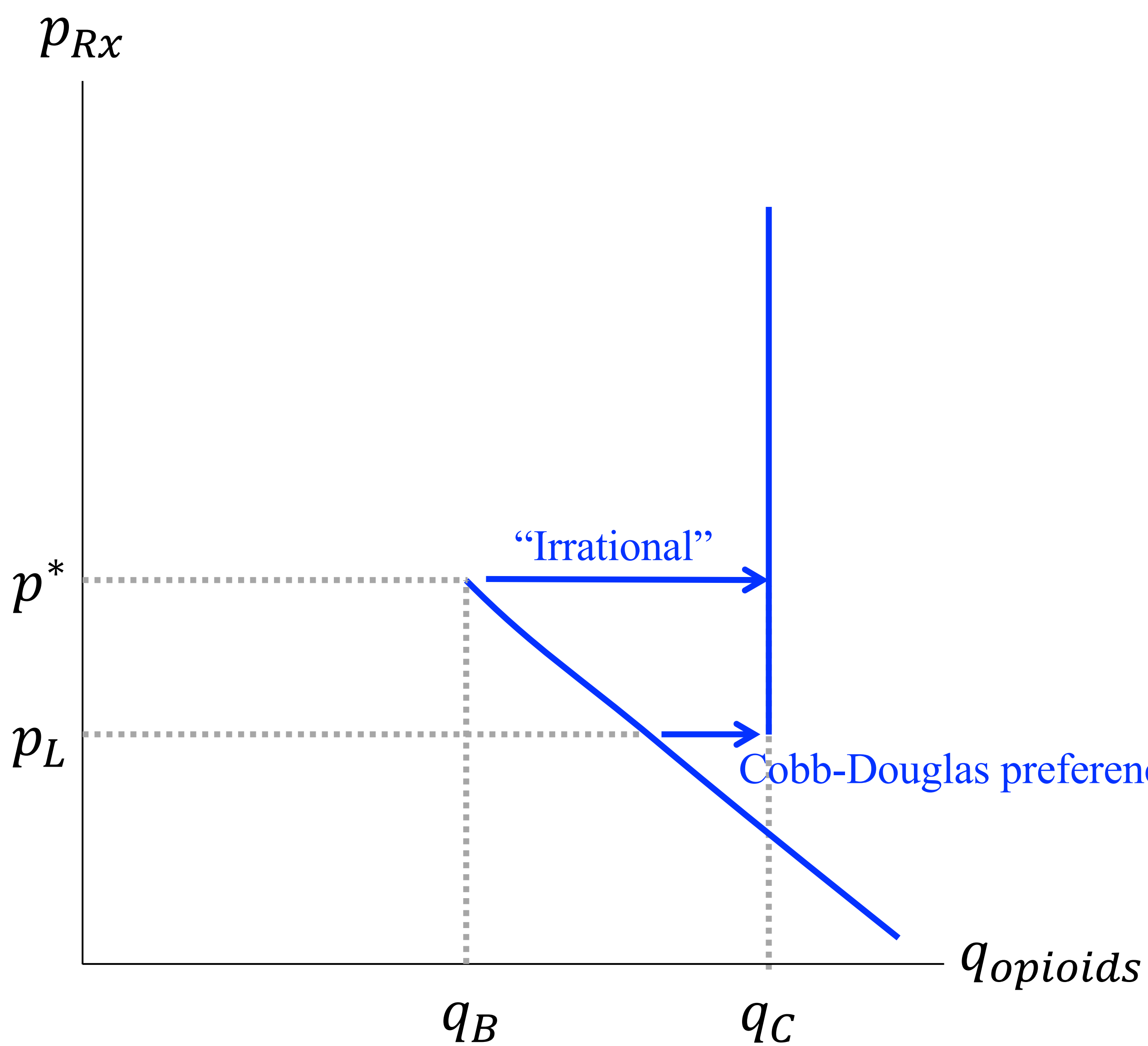


Figure 11. Rational and irrational demand curves as a function of the Rx price.

(Price and quantity are shown on log scales)



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