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Urban Water Disinfection and Mortality Decline in Lower-Income Countries

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Historically, improvements in municipal water quality led to substantial mortality decline in today's wealthy countries. However, water disinfection has not consistently produced large benefits in lower-income countries. We study this issue by analyzing a large-scale municipal water disinfection program in Mexico that increased water chlorination coverage in urban areas from 58% to over 90% within 18 months. We estimate that the program reduced childhood diarrheal disease mortality rates by 45 to 67%. However, inadequate sanitation infrastructure and age (degradation) of water pipes may have attenuated these benefits substantially.

JEL: I10, I18, O18

Historically, improvements in the quality of municipal drinking water made important contributions to population health in today's wealthy countries. Late nineteenth and early twentieth century investments in water purification led to substantial reductions in urban mortality in a number of countries, including Japan, France, Sweden, the United States, and the United Kingdom – in some cases, virtually eliminating waterborne disease (Cain and Rotella, 2001, Cutler and Miller, 2005, Ferrie and Troesken, 2008, Ketzenbaum and Rosenthal, 2014, Knutsson, 2016, Koppaka, 2011, Ogasawara and Inoue, 2015, Preston and van de Walle, 1978). These disinfection technologies were often

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introduced into relatively new municipal water systems with good quality pipes and sufficient supply to deliver water at full pressure without interruption (Melosi, 2000).

Despite century-old knowledge about the benefits of water disinfection (Turneure and Russell, 1901), diarrheal diseases due to poor water quality remain prevalent in many low- and middle-income countries today. Worldwide, diarrhea remains a leading cause of child mortality (Liu et al., 2012, Troeger et al., 2018, Vos et al., 2015) as well as a leading cause of morbidity (roughly 1.7 billion episodes per year) (Fischer Walker et al., 2013b). This disease burden persists despite the fact that cities in developing countries are using many of the same disinfection technologies that were historically successful (Bain et al., 2014, Ercumen et al., 2015, Gadgil, 1998).

This study assesses the effectiveness of water disinfection programs in lower-income countries, focusing on the case of Mexico. In doing so, it also seeks to identify factors that limit the realization of their full potential. Recent work on improving water quality in low- and middle-income countries has generally focused on interventions that are either small in scale or target rural populations with limited water system infrastructure (Arnold and Colford, 2007, Arnold et al., 2013, Clasen and Haller, 2008, Duflo et al., 2015, Gruber et al., 2013, Kremer et al., 2011, Zwane and Kremer, 2007). However, the urban population is projected to reach nearly 2.5 billion worldwide by 2050 (United Nations Population Division, 2014), and the rapid growth of cities will require larger-scale interventions focused on municipal water supply (Brown et al., 2009, Fotso et al., 2007, McDonald et al., 2014, United Nations Development Programme (UNDP), 2006). The two studies (of which we are aware) examining interventions to improve urban water quality on a large scale are Galiani, Gertler, and Schargrodsky (2005), who show that privatization of municipal water companies in Argentina led to significant reductions in child mortality from water-borne diseases, and Greenstone and Hanna (2014), who find no effects of environmental regulation on water pollution in Indian cities. Rather than focus on ownership or regulation, we focus specifically on disinfection technology, which is a cornerstone of urban water provision.

Improving urban water quality in developing countries on a large scale with otherwise proven disinfection technologies is difficult for many interrelated reasons. First, municipal water systems in many low- and middle-income countries are older, and

thus subject to degradation. Pipe breaks and breaches allow even clean water to become contaminated (or re-contaminated) by bacterial pathogens in the surrounding soil (Lee and Schwab, 2005, Shaheed et al., 2014). Low or intermittent water pressure in pipe networks can further reduce the efficacy of chlorination by increasing the degree of re-contamination (Bhutta et al., 2013, Ercumen, et al., 2015, Jeandron et al., 2015, Kumpel and Nelson, 2013, Tokajian and Hashwa, 2003).¹ Both problems are compounded by the fact that repairing degraded infrastructure (or building new distribution networks) may be prohibitively expensive, and financing mechanisms (either through domestic capital markets or partnerships with multinational organizations) may be unavailable (Crocker and Masten, 2002, Cutler and Miller, 2006, Gadgil, 1998, Masten, 2011, United Nations Development Programme (UNDP), 2006). Second, improvements in water quality may not be effective without complementary investments in sanitation (Alsan and Goldin, 2019, Duflo, et al., 2015). Third, the provision of improved water may weaken private incentives for protective health behaviors. Such compensatory behavior holds the potential to crowd-out the health benefits of water disinfection (Bennett, 2012, Keskin et al., 2015).

A central contribution of our paper is that we study a massive nation-wide program to chlorinate municipal water systems across the entire country of Mexico within a single year. Because conditions across Mexico's municipalities varied widely at the time of implementation, we are able to investigate factors governing program success (and failure) in improving health. *Programa Agua Limpia* (henceforth, PAL), a major clean water program, was launched in 1991 in response to a cholera epidemic that swept rapidly through Central and South America (Gutierrez et al., 1996, Sepulveda et al., 2007, Sepulveda et al., 2006b).² Within 6 months, the share of Mexico's urban population receiving disinfected water rose dramatically from 58% to over 90% (CONAGUA,

¹ Unlike other disinfection technologies (like filtration, for example), residual chlorine remains in water from point of treatment to point of consumption. However, low or intermittent water pressure (on which data are not systematically available) increases the risk of recontamination as stagnant organic matter in the water supply effectively absorbs and reduces circulating chlorine levels. We thank Steve Luby for alerting us to this point.

² Notably, the baseline burden of diarrheal disease in Mexico around this time rivaled that of modern low-income countries. For example, data from the Global Burden of Disease project demonstrate that there were 900 deaths per 100,000 post-neonates (1-12-month-olds) in Mexico in 1990. In low-income countries as a whole the corresponding rate in 1990 was 1,300 per 100,000 and 500 per 100,000 in 2015. See <http://ghdx.healthdata.org/gbd-results-tool>.

1994), with continued increases through 1994 (The World Bank, 1995). The majority of this growth occurred in small and medium sized towns and cities, as larger cities had already chlorinated their water supplies prior to PAL. Importantly, clean water coverage was achieved without significant expansion of existing piped water infrastructure or sewage networks.

To estimate the impact of PAL on child mortality across Mexico, we use detailed, cause-specific administrative mortality statistics at the municipality level. While there was little variation in the timing of PAL implementation across regions – and there is no available data on exact program activities at the municipal or state levels – we are able to distinguish program effects on diarrheal deaths from potentially correlated omitted trends by using multiple difference-in-difference approaches. Specifically, we assess changes in child diarrheal mortality rates in small and medium sized municipalities against childhood diseases not directly influenced by water quality as well as against changes in diarrheal mortality rates in municipalities containing (or within) large cities with 500,000 or more residents, whose water supplies were already chlorinated prior to PAL.³ We also examine program effects on all-cause child mortality, using larger cities as control cases. These complementary approaches collectively account for several potential threats to inference, including program spillover effects on other causes of death and improvements in cause of death registration during the study period. In all specifications, we allow for differential pre-existing trends across treatment and comparison groups, accounting for any decline in diarrheal disease mortality due to either secular improvements of living standards or prior public policy responses (for example, large-scale distribution of oral rehydration salts (ORS) in 1985 (Frenk et al., 2003, Sepulveda et al., 2006a, Sepulveda, et al., 2007).

These different approaches yield substantively similar results. Our estimates suggest that PAL was associated with a 45%-67% relative reduction in diarrheal disease mortality rates among children under age 5, averting 9,100-16,000 deaths over the period

³ The use of larger cities as controls helps account for (minor) cholera prevention and treatment activities other than chlorination that were common across smaller towns and large cities – including legal restrictions on wastewater irrigation, hygiene campaigns, provision of Oral Rehydration Therapy (ORT), and general disease surveillance (CONAGUA, 1994, Sepulveda, et al., 2006b). In practice water system chlorination was the overwhelming focus of the program (CONAGUA, 1994).

of 1991-1995 alone. We conservatively estimate cost per life year saved by PAL of \$1,896-3,392 (2015 USD), suggesting that the program was cost-effective.⁴ Our substantive findings are robust to using different independent variable transformations and as well as death counts instead of death rates, and are unlikely to be driven by differential migration.

Motivated by the contrasting experience of wealthy countries in history and lower-income countries today, we then study municipal-level variation in program effects to investigate the circumstances under which large-scale municipal water disinfection in developing countries can be successful. First, we find that reductions in diarrheal disease mortality rates were not larger in urban municipalities with greater baseline piped water coverage rates, despite the fact that piped water infrastructure was required to disseminate chlorination. Because more extensive systems in Mexico are generally older, this result is consistent with engineering concerns about recontamination through degraded, aged infrastructure (Lee and Schwab, 2005, Mazari-Hiriart et al., 2005, Tulchinsky et al., 2000) and reduced chlorine efficiency due to irregular water pressure and outages (Ashraf, et al., 2017, Ercumen, et al., 2015, Kumpel and Nelson, 2013, Kumpel and Nelson, 2014), and we provide suggestive evidence in support of this hypothesis.

Second, we find that the health benefits of clean piped water were greater in municipalities with more extensive sewage infrastructure. This finding is consistent with emerging evidence of complementarities between water disinfection and sanitation in the economics literature (Alsan and Goldin, 2019, Duflo, et al., 2015), although it stands in contrast with an epidemiology literature suggesting no such complementarities (among more localized interventions Fewtrell et al. (2005)).

Notably, predictions using our estimates suggest that childhood diarrheal disease mortality rates would have declined by 90% in urban areas nationwide if all municipalities had the same degree of sewage infrastructure as those in top quartile. This

⁴ These estimates likely underestimate the true cost-effectiveness of PAL for two reasons. First, due to limitations in available cost data, we are only able to consider the period 1991-1995, but the health benefits of PAL presumably accrued far beyond this period. Second, recent studies show that clean water may have important effects on time allocation, financial transactions, and long-run human capital accumulation (Ashraf et al., 2017, Bhalotra and Venkataramani, 2015), but our calculation focuses entirely on survival benefits.

degree of mortality decline is broadly consistent with the historical health benefits of municipal water disinfection in today's wealthy countries. By contrast, areas with the lowest sewage coverage experienced considerably attenuated reductions in diarrheal disease mortality rates under PAL.⁵

Our results have two broad policy implications. First, the average effects that we estimate for childhood waterborne deaths are large, suggesting that despite misaligned incentives and political manipulation that plague many state programs (Galiani et al., 2005, Greenstone and Hanna, 2014), there is nonetheless the potential for state-led initiatives to have substantial health impact.⁶ Second, our finding that the effectiveness of urban water disinfection depends heavily on infrastructure quality, which is generally poor in lower-income countries, suggests a potential role for improvements in financing and access to capital – through private-public partnerships or innovations in local public finance, for example (Crocker and Masten, 2002, Cutler and Miller, 2006).

The rest of this paper is organized as follows. Section 2 provides background on Mexico's clean water reform, *Programa Agua Limpia*, and Section 3 describes our data. Section 4 presents changes in waterborne disease death rates associated with PAL, Section 5 examines the circumstances under which public water quality investments are most effective, Section 6 presents estimates of the cost-effectiveness of PAL, and Section 7 concludes.

I. The History of *Programa Agua Limpia* in Mexico

As in many developing countries, infectious diseases have historically been responsible for most of Mexico's burden of disease among children. During the 1980s, diarrheal diseases and acute respiratory infections were the two leading causes of child death, and diarrhea alone was responsible for nearly a quarter of deaths under age 5 (Gutierrez, et al., 1996). Diarrheal mortality was concentrated in the poorer southern

⁵ In an earlier working paper, we also examined behavioral responses, finding that households in states benefitting more from PAL (on the basis of pre-intervention diarrheal disease deaths) reduced their spending on bottled water, soaps, and detergents. Given the large impact on mortality that we find, we concluded that crowd-out of protective private health behaviors was far from complete.

⁶ In fact, the average treatment effects we estimate are larger than found in Galiani, Gertler, and Schargrodsky (2005), who find that privatization led to an 8% decrease in child diarrheal disease mortality on average (26% in the poorest areas).

region of Mexico (*Figure 1*), with rates in small- and medium-sized towns nearly twice as high as those in cities.⁷ Mexico's primary approach to controlling diarrheal diseases over this period emphasized expansion of access to Oral Rehydration Therapy and targeted clinical case management (Frenk, et al., 2003, Gutierrez, et al., 1996, Mota-Hernandez and Velasquez-Jones, 1985, Sepulveda, et al., 2007, Sepulveda, et al., 2006b).

In 1991, a cholera pandemic emerged in Chile and Peru and quickly spread through South and Central America (Medina, 1991, Ries et al., 1992, Sepulveda, et al., 2006b). In an effort to limit its spread across Mexico, the Mexican Ministry of Health and the newly created National Water Commission (*Comision Nacional del Agua*, CONAGUA) launched *Programa Agua Limpia* (PAL, or the National Clean Water Program) in April 1991 (Sepulveda, et al., 2007, Sepulveda, et al., 2006b). Thus, the introduction of the program was generally unanticipated and driven by external factors. PAL was in principle a multi-faceted campaign that included: (1) chlorination of previously untreated water sources; (2) restrictions on the use of wastewater for irrigation (an important component of Chile's efforts against cholera (Medina, 1991)); (3) health education campaigns targeting both the general population and health care providers; and (4) expansion of the availability of Oral Rehydration Therapy (ORT).

In practice, chlorination of municipal drinking water was the centerpiece of the program (CONAGUA, 1994, World Bank, 1994). Because larger cities (those with more than 500,000 inhabitants) already had established chlorination and filtration systems, PAL chlorination efforts targeted small and medium sized cities. In contrast, other (minor) program components were not specifically targeted to these areas and implemented more broadly (Sepulveda, et al., 2006b). Within 6 months of PAL's launch (*Figure 2*), access to disinfected water increased from 39 million to 61 million individuals covered (CONAGUA, 1994).⁸ This corresponds to an increase in disinfection coverage from 58% to well over 90% in the urban population in Mexico (and from 42% to 72% for the population as a whole).

⁷ Averaging over 1985-1990, the child diarrheal mortality rate in municipalities predominantly containing small towns and medium-sized cities was 7.76 deaths per 1,000 live births. In large cities (>1 million population), the corresponding mortality rate was 4.48 per 1,000 live births.

⁸ An additional 8 million individuals were covered between 1991 and 1994, thus leading to near universal disinfection coverage in the urban population.

Achieving these gains required a vast expansion of municipal water disinfection predominantly through *existing* water pipe infrastructure. Nationwide, the number of water treatment plants grew from 250 to nearly 15,000, the number of residual chlorine monitoring stations expanded from 200 to over 100,000 sites, and treatment capacity (the volume of chlorinated water per unit time) doubled. In addition, to improve water quality in areas without piped water coverage, chloride tablets were disseminated to households (Sepulveda, et al., 2006b) and quality monitoring for commercial bottled water and ice was expanded (CONAGUA, 1994), although these efforts were secondary relative to disinfection of water at the source.⁹

Previous work on PAL hints at potentially large program effects on health. Gutierrez et al. (1996), Velazquez et al. (2004)), and Sepulveda et al. (2006) show sharp declines in childhood diarrheal disease mortality rates beginning in 1991. Focusing on morbidity, Gutierrez et al. (1996) analyze changes in aggregate morbidity rates, showing that the average number of annual episodes of diarrheal disease morbidity among children decreased from 4.6 to 2.2 between 1990 and 1993. Similarly, Velazquez et al.

⁹ Similarly, other components of PAL received substantially less emphasis and are also less likely to have had a meaningful impact on diarrheal disease mortality. For example, legal restrictions reduced the amount of farmland irrigated with wastewater (World Bank, 1994). However, only a very small fraction of farmland was irrigated with wastewater prior to PAL (~0.1%) (CONAGUA, 1994). Even if these crops were widely distributed, they are likely to have been in large cities as well as in small and medium-sized towns. Thus, we are able to account for this component of the policy in our triple difference models – see Section 4.1. Efforts to modify hygienic behavior included use of radio and television to disseminate messages about hand-washing and other sanitary practices (Sepulveda, et al., 2006b). However, experiences from other low- and middle-income settings suggest that their impact on diarrheal disease mortality is likely to be small, particularly in comparison with water system disinfection (Ahuja et al., 2010, Dupas, 2011). In addition to PAL-related efforts, two other potentially relevant programs are Programa Nacional de Solidaridad (PRONASOL), a nationwide anti-poverty program implemented between 1989-1994, and Programa Nacional de Agua Potable, Alcantarillado y Saneamiento (PAPAZU), for which Mexico received a \$300 million loan from the World Bank to expand access to potable water, sanitation, and sewage in the early 1990s (The World Bank, 1995). However, these programs are unlikely to confound our analysis. PRONASOL-related potable water investments were generally focused in rural areas (Diaz-Cayeros and Magaloni, 2003), which we exclude from our analysis. PAPAZU was small in scope: the share of the population covered by clean water and sanitation services rose much less than the share covered by disinfection. Put differently, the 3.2 million individuals benefitting from PAPAZU (The World Bank, 1995) represents only a fraction of the 30 million individuals who gained access to chlorinated water in the first year of PAL. PAPAZU activities were also concentrated in states (Estado de Mexico, Guanajuato, and Sonora) that had relatively lower pre-program burdens of diarrheal disease. Removing these states in our main regression models does not change our results. Finally, PAPAZU project reports reveal that many of the projects initiated as part of PAPAZU were deemed “out of compliance,” and thus potentially of lower quality (The World Bank, 1995).

(2004) show that morbidity declined by over 63% during the period 1990-1995. However, these studies essentially only describe national trends over time.

II. Data and Descriptive Evidence

A. Data Sources

For our primary analyses, we use data from two main sources: the Mexican Vital Statistics and a 10% sample of the 1990 Mexican Population Census.¹⁰

Mexican Vital Statistics. We use mortality data for infants and children under 5 years old from the Mexican Ministry of Health. The vital statistics contain individual-level records of every certified death in the country.¹¹ Each death record contains information about the cause of death (coded using the International Classification of Diseases, 9th Edition, or ICD-9), age at death, and municipality of death. Municipalities are the next administrative division below the state (analogous to U.S. counties). ICD-9 coded data are available for the period 1979-1997 (thereafter ICD-10 codes were used to delineate causes of death).

We aggregate individual-level deaths for children under the age of 5 into municipality-cause-year cells. Given our focus on PAL, we create a category for infectious diarrheal diseases using three-digit codes from the International Classification of Diseases, Ninth Revision (ICD-9).¹² To convert diarrheal death counts into mortality rates, we use data on the number of live births in each municipality and year.¹³

In our main empirical models, we restrict our analysis to the period 1985-1995. The starting time point reflects the period at which municipality-year data on live births

¹⁰ Data and code are deposited at ICPSR (Bhalotra et al., 2021).

¹¹ Certified deaths are all deaths that are brought to the attention of the National Statistics Office, Health Ministry, Judiciary, Military, and funeral directors. Certification, which includes coding of cause of death, is provided by an individual with a license to do so (typically a physician, nurse, or Health Ministry representative).

¹² We used ICD-9 codes 001-009 to identify deaths from infectious diarrhea. These codes cover intestinal infectious diseases and include cholera (001), typhoid and paratyphoid (002), salmonella infections (003), shigellosis (004), food poisoning (005), amoebiasis (006), protozoan causes (such as giardia and cryptosporidiosis, 007), infections due to other organisms (such as rotavirus and other viruses, 008), and presumed intestinal infection due to an ill-defined cause.

¹³ We obtained municipality-year data on live births from the *Instituto Nacional de Estadística y Geografía* (INEGI <http://en.www.inegi.org.mx/proyectos/registros/vitales/natalidad/>). To compute mortality rates, we used a direct method (UNICEF et al., 2007) and divided the number of child deaths for a given cause-municipality-year by the number of live births in the same municipality-year (scaled to reflect the number of deaths per 1,000 live births).

are available. The latter time point marks the end of PAL program efforts. We do, however, use data on the number of diarrheal deaths between 1979-1997 in descriptive analyses identifying structural breaks (for which a longer time series is required for statistical power).

A note about the quality of Mexico's vital statistics is warranted. As in many developing countries, there are important concerns about under-reporting of deaths, particularly in poorer regions and among young children (Hernandez et al., 2012, Lozano-Ascencio, 2008, Tome et al., 1997).¹⁴ Corrections for under-reporting and misclassification of deaths were made by the Ministry of Health from 1980 onwards, and these data have been used in other prominent studies, albeit at levels of aggregation above the municipality level (Barham, 2011, Cutler et al., 2002, Foster et al., 2009, Gonzalez and Quast, 2011). In part because of these efforts, Mexico's Vital Registration system is now considered one of the best in the developing world in terms of completeness and quality (Mathers et al., 2005).

Nevertheless, we cannot rule out the possibility that some degree of underreporting remains. However, for this to bias our results, underreporting must have changed sharply in 1991, and either differentially across smaller- vs. larger-municipalities or for diarrheal diseases relative to control diseases.¹⁵ We investigated changes in the *quality* of death records – measured by missing or “unknown” causes of death – and concluded that these cases are too few to bias our findings. Specifically, for the period 1985-1997, we assessed the quality death records on two margins. First, we examined the number of death records for which the cause of death was missing or not coded, finding that 0% of death records had missing causes. Second, we assessed the prevalence of causes of death coded as “ill-defined or unknown causes of mortality.” While the share

¹⁴ Hernandez et al. (2012) show that infant and child deaths may be underreported by as much as 20% in a sample of low human development index municipalities in 2008. However, it is important to note that in many instances birth certification might also underreport those children. In this case, the counts in the numerator and the denominator of reported mortality rates are under-counts and, in general, it is difficult to sign any resulting bias in the rates.

¹⁵ The direction of any bias in the estimated treatment effect is *a priori* unclear. If improvements in measurement occurred across the board, then our use of differences with respect to control diseases will help account for bias from measurement error. However, it may be that recording of diarrheal disease mortality in particular improved with the introduction of PAL (via investments in diarrheal disease recognition and surveillance). This would bias downward our estimates of mortality rate changes associated with PAL.

does decline slightly after 1990, the prevalence of unknown cause codes is less than 1% through the period (*Figure A1*). In our empirical analysis, we nonetheless address the possibility that ascertainment and registration of causes of death improved over the study period by examining all-cause child mortality – which would not depend on the quality of cause of death coding - as an outcome.

1990 Mexican Population Census. We use the population census (a) to identify relevant municipalities (i.e., those targeted by PAL or those used for comparison) and (b) to obtain municipal characteristics for estimating heterogeneous PAL program effects. Using individual-level information on urban versus rural residence in a 10% sample of Mexico’s 1990 population census (Minnesota Population Center, 2015), we compute the proportion of each municipality’s population living in small- and medium-sized cities versus large cities (defined by the Mexican national statistical agency as having 500,000 inhabitants) and rural areas (fewer than 2,500 inhabitants).¹⁶ The former distinction allows us to distinguish between the small-and medium-sized urban areas that were exposed to PAL chlorination efforts against big cities, which already had chlorination infrastructure in place well before PAL. We exclude from our sample fully rural municipalities as they, too, were not the primary focus of PAL chlorination efforts, though were targeted by other policies (see Section 2.1).

To analyze heterogeneous effects of PAL across municipalities, we use household-level data from the 1990 census to create aggregated municipal-level measures of the share of households with piped water and the share of households with sewage system connections in their dwelling (a measure of sanitation coverage). Additionally, we create measures of average earned income among adults; the fraction of adults completing secondary schooling; and the fraction of the population speaking an indigenous language. We use the 1960 and 1990 census microdata to compute measures

¹⁶ An individual’s town of residence in the 1990 census was divided into population sizes of 1-2,499, 2,500-14,999, 15,000-99,999, 100,000-499,999, and 500,000 and above. We calculated the fraction of individuals living in towns of 500,000 or more in each municipality to capture large cities and the fraction living in towns of 1-2,499 inhabitants to capture rural areas. Any municipality that had a non-zero fraction living in towns of 500,000 or more was coded as a large city, and the small area dummy is simply the inverse. We removed municipalities in which 100% of respondents lived in rural areas, which applied to 970 of 2,394 municipalities in our data.

of piped water system age. *Tables 1 and 2* show descriptive statistics for all variables used in our analyses.

B. Descriptive Evidence

Figure 3 plots trends in the natural logarithm of the child (under age 5) death rate from diarrheal diseases between 1979-1997 at the national level, along with the predicted rates that would arise if diarrheal disease mortality had followed the same trend as between 1979-1985.¹⁷ The series displays a steady decline in diarrheal deaths from 1979-1990, followed by a trend break in 1991. We formally tested for a structural break between 1985-1995 (Quandt, 1960), remaining agnostic about the exact break point in the time series, and calculate the F -statistic on different user-specific break points in the window. The largest F -statistic across tests of different time points is used to identify the break point. For diarrheal diseases, examining each year between 1985-1995 as potential break points, the test identifies a break in 1991, coinciding with the timing of PAL ($F = 7.82$, $p = 0.013$).¹⁸

Figure 4 previews our difference-in-differences strategy, comparing trends in (1) child mortality (per 1,000 live births) from diarrheal vs. two sets of “control” diseases - respiratory infections and non-infectious childhood conditions - over 1985-1995, the tight window we use in our primary analyses (see Section 3.1); and (2) child diarrheal disease and all-cause mortality rates across small- and medium-sized cities (targeted by PAL chlorination efforts) vs. control large cities (not targeted by virtue of pre-existing disinfection infrastructure) over the same time frame. The motivation for the use of each set of controls is provided in Section 4.1.

Among small-and medium-sized cities, mortality from diarrhea – the leading cause of child death prior to PAL - dropped below mortality rates from each of the control diseases by 1992 (*Figure 4*, Panel A). While diarrheal disease mortality was already declining relative to control diseases before PAL (a fact we explicitly address in our empirical models), the sharp trend break seen for diarrheal diseases was not observed

¹⁷ We chose the period 1979-1985 as (1) this was prior to any large-scale efforts to reduced diarrheal disease deaths and (2) to not overlap with the period of time (1985-1995) over which we search for structural breaks. In practice, fitting a trend at any time point prior to 1991 produced similar findings.

¹⁸ We ran this test after detrending the time series using the estimated linear time trend over 1979-1985, after which the null hypothesis of a unit root in the time series prior to the implementation of PAL was rejected by an augmented Dickey-Fuller test. See Hansen (2001)).

for the control diseases.¹⁹ In addition, between 1991 and 1995, child diarrheal mortality rates – as well as all-cause mortality rates - in small- and medium-sized cities converged to those in larger, previously chlorinated cities (*Figure 4*, Panels B and C), with pre-1991 trends in mortality rates similar across the two groups.

III. Water Disinfection and Diarrheal Disease Mortality Rates

A. Empirical Strategy

To formally estimate the relationship between PAL and mortality rates among children under-5, we employ two complementary sets of difference-in-differences strategies. First, we test for differential trend breaks in diarrheal disease mortality in small and medium-sized cities relative to two sets of control diseases which should not be directly affected by PAL.²⁰ These specifications allow us to address the possibility of overall improvements in the disease environment that may have led to convergence between diarrheal mortality rates in small-and medium-sized versus large cities (though such omitted factors are less likely given that we have established a coincidence in timing of the trend break in diarrheal mortality and the introduction of PAL program efforts). One group of control diseases is acute upper and lower respiratory infections, including viral bronchitis and pneumonia. We choose these diseases because they were the second-leading cause of child mortality in Mexico prior to PAL²¹ and because they share several common risk factors with diarrheal diseases (respiratory infections are spread through oral droplets and diarrheal diseases are spread by fecal-oral contamination). Particularly helpful for identification is the fact that respiratory diseases are more sensitive to water *quantity* (which affords opportunities for preventive handwashing and thus breaking the droplet contamination cycle), while diarrheal diseases are more sensitive to water *quality*, which is the focus of PAL (Ahuja, et al., 2010). That is, poor drinking water quality is not

¹⁹ Using (detrended) time series data on logged deaths for 1979-1997 (see *Figure A2*), and again searching for breaks over 1985-1995, we do not find evidence of any trend breaks timed with PAL for our “control” diseases, namely respiratory diseases (break year 1987, $F = 2.76$, $p=0.12$) or non-infectious childhood diseases (break year 1989, $F = 6.9$, $p=0.018$).

²⁰ A similar strategy was employed by Jayachandran et al. (2010) in their work on antibiotic therapy in the United States. Closer to our study, Galiani, et al. (2005) use non-diarrheal causes of mortality as a falsification test in their study of water service privatization in Argentina.

²¹ ICD-9 codes 460-466 and 480-487, respectively.

a *direct* risk factor for respiratory infections (Fischer Walker, et al., 2013b).

However, diarrheal diseases weaken immune systems and may therefore increase the likelihood of contracting other infectious diseases (Fischer Walker et al., 2013a, Sedgwick and MacNutt, 1910). Historical work suggests that respiratory diseases in particular may fall with changes in water quality – the so-called Mills-Reincke phenomenon (Alsan and Goldin, 2019, Cutler and Miller, 2005, Ferrie and Troesken, 2008, Inoue and Ogaswara, 2020). For this reason, we also use a second, alternative control group of diseases for which the Mills-Reincke effect is likely less salient (i.e., that are biomedically more ‘distant’). To construct this group of diseases, we focus on deaths from perinatal causes (low birth weight, birth trauma, congenital infections) and congenital anomalies for which poor water quality is not a known direct risk factor (Embrey et al., 2004, Pruss-Ustun et al., 2008).²²

Second, we test for differential trend breaks in diarrheal disease mortality rates in 1991 (when PAL was implemented) across small- and medium-sized municipalities, which were targeted by PAL chlorination efforts, versus large cities, which already had long-standing chlorination infrastructure prior to PAL. This comparison allows us to both address any residual concerns from cross-disease spillovers²³ and net out any effects of other (minor) components of PAL - such as legal restrictions on wastewater irrigation, hygiene campaigns, some provision of Oral Rehydration Therapy (ORT), and general disease surveillance (CONAGUA, 1994, Sepulveda, et al., 2006b) – though we anticipate these effects would be small.

A residual concern across all of these strategies is endogenous change in the quality of cause of death ascertainment. For example, given the PAL was implemented due to concerns about cholera, it is possible that ascertainment of diarrheal disease deaths improved over the study period as authorities sought to better track diarrheal diseases. To address this concern, we test for differential trend breaks in child mortality from all-causes across small-and medium versus large municipalities.

²² ICD-9 codes are as follows: congenital anomalies (ICD9 codes 740-759); perinatal causes (low birth weight, birth trauma, congenital infections, neonatal jaundice, etc: ICD9 codes 764-779).

²³ Co-dependence between diarrheal diseases and these more bio-medically distant controls is still possible if, for example, falling diarrheal disease mortality raises the returns to health behaviors and investments more generally, which would potentially reduce the risk of all of the control diseases (Dow et al., 1999).

We view these strategies as complementary approaches. Comparisons with control diseases allow us to better account for changes in shared risk factors, but are more prone to (downward) bias from cross-disease spillovers. To internalize these spillovers, we also present estimates for all-cause mortality, which addresses any potential endogenous changes in the accuracy of cause of death reporting. By contrast, comparisons across small-medium versus large cities are less prone to these spillovers. However, epidemiologic trends may differ across large versus small-cities (Gomez-Dantes et al., 2016), introducing bias into these comparisons.²⁴ In addition, the sample size for the control group is small: there are only 16 municipalities containing large cities. For these reasons, we view the specifications using respiratory diseases as controls as the most conservative.

Formally, for each of our comparisons, we begin by estimating versions of the following event study specifications:

$$(1) Mort_{djt} = \sum_{t=1985}^{1995} \delta_t (1(Diarrhea_d) \times 1(Year = t)) + \sum_{t=1985}^{1995} \mu_t(1(Year = t)) + \alpha_d + \lambda_j + u_{djt}$$

$$(2) Mort_{jt} = \sum_{t=1985}^{1995} \alpha_t (1(Small_j) \times 1(Year = t)) + \sum_{t=1985}^{1995} \mu_t(1(Year = t)) + \lambda_j + e_{jt}$$

In models using respiratory and non-infectious childhood diseases as controls (Equation (1)), we restrict our sample to small-and medium-sized municipalities. In these models, $Mort_{djt}$ represents the inverse hyperbolic sine transformation²⁵ of the mortality

²⁴ An additional concern is that using already treated areas as controls may be biased treatment effects downwards (Callaway and Sant’Anna, Forthcoming, Goodman-Bacon, 2018). However, this possibility is less likely in the case of PAL, as disinfection in larger cities occurred well before our study period. We also explicitly net out area-specific pre-trends in our modelling approach (see Equation 4). Lastly, this potential bias is more likely to materialize when the timing of treatment varies across units. This does not apply to the case of PAL, where the disinfection treatment was effectively rolled out within a single calendar year.

²⁵ Formally, the inverse hyperbolic sine transform of a variable, y , is $\ln(y + \sqrt{y^2 + 1})$. The advantage of this transformation is that it is defined at zero (Burbridge et al., 1988). Estimates obtained using this transformation can be interpreted in the same manner as those obtained using a natural logarithm transformation of the dependent variable if there is a relatively small number of zeros (per a recent rule of thumb, less than 1/3 of all observations). See Bellemare and Wichman (2020) for further details. In our case, the share of zeros ranges from 28% to 31%, depending on the specific double- or triple-difference sample used. Nevertheless, we explore other econometric methods: our substantive estimates are robust to

rate for disease class d in municipality j and year t and $1(Diarrhea_d)$ denotes whether or not the cause of death is diarrheal disease versus either respiratory or non-infectious childhood control diseases. In models using large cities as controls (Equation (2)), $Mort_{jt}$ represents either the inverse hyperbolic sine transformation of the diarrheal mortality rate or the overall child mortality rate in municipality j and year t and $1(Small_j)$ is a dummy variable equal to 1 for the targets of PAL chlorination efforts - small- and medium-sized municipalities - versus large cities. In both sets of models, μ_t capture secular trends in non-treated municipalities or control diseases, $1(Year = t)$ denotes observations in year t , and λ_j represent municipality fixed effects.

The parameters of interest are captured by the vector δ_t and α_t , which respectively yield the average differential percentage change in mortality rates from diarrheal diseases relative to mortality rates for control diseases and diarrheal and all-cause mortality rates in small- and medium-sized cities versus large cities in year t relative to the baseline year (which we set as 1990). We estimate a set of lags and leads and plot the coefficients with confidence intervals.

Then, to estimate average program effects of PAL during the study period (accounting for any differential pre-trends by disease or city size), we use parametric specifications of the following form:

$$(3) \text{Mort}_{djt} = \theta_0 + \theta_1 (1(Diarrhea_d) \times 1(Post_t)) + \theta_2 (1(Diarrhea_d) \times 1(Post_t) \times Year_t) + \theta_3 (1(Diarrhea_d) \times Year_t) + \theta_4 1(Diarrhea_d) + \sum_{t=1985}^{1995} \mu_t (1(Year = t)) + \lambda_j + u_{djt}$$

$$(4) \text{Mort}_{jt} = \beta_0 + \beta_1 (1(Small_j) \times 1(Post_t)) + \beta_2 (1(Small_j) \times 1(Post_t) \times Year_t) + \beta_3 (1(Small_j) \times Year_t) + \sum_{t=1985}^{1995} \mu_t (1(Year = t)) + \lambda_j + e_{djt}$$

Here, $1(Post)$ is a dummy variable for post-PAL years (1991 and later), and all other variables are defined as before. The parameters of interest are β_1 and β_2 (Equation (3)) and θ_1 and θ_2 (Equation (4)), which measure level and trend breaks beginning in 1991 for diarrheal disease mortality rates relative to control disease mortality rates. Importantly,

other transformations that preserve observations with zeros, such as a quartic root transformation, as well as count data methods.

the trend and level breaks are net of any pre-existing differential trends in treatment and control areas or conditions (captured by β_3 and θ_3 , respectively). Across all specifications, we consistently cluster the standard-errors at the municipality level.

We assess the robustness of our substantive findings to alternate transformations of the dependent variable, specifically the linear transform and the quartic root transform. We also estimate negative binomial versions of our models using death counts instead of rates. Finally, we assess the robustness of our standard errors to clustering at the state (instead of municipality) level.

B. Results

Figure 5 plots event study estimates obtained by estimating Equations (1) and (2). Across models – and consistent with the unadjusted graphs (*Figure 4*) – the coefficient series show a discrete, rapid decline in treatment relative to control group estimates, coinciding with the implementation of PAL in 1991 and continuing throughout the post-implementation years. The presence of modest pre-existing trends, particularly in the case of the respiratory disease controls, highlights the importance of empirically accounting for disease-specific mortality rate trends.

Table 3 then reports estimates from Equation (3) and (4), which explicitly adjust for differential pre-existing trends. We find statistically significant breaks in diarrheal disease mortality rates to both sets of control diseases and untreated large cities, and these breaks coincide with the introduction of PAL. Specifically, in models using control diseases, we find both statistically significant level and trend breaks (θ_1 and θ_2 in Equation (3), respectively), on top of negative pre-trends (θ_3). In models comparing diarrheal disease mortality rates across small and large cities, we find a statistically significant trend break (β_2 in Equation (4)) relative to flat pre-existing trends (near zero point estimates for β_3) for both diarrheal disease and all-cause mortality. In all cases, the estimate of the trend break alone is nearly three or more times as large as the estimated pre-trends.

Collectively, the different models yield similar substantive findings: the estimates imply a 45-67% reduction in diarrheal disease mortality rates relative to control diseases or control cities by 1995, implying 9,149-16,437 deaths averted between 1991-1995 as a

result of PAL.²⁶ The smallest estimates are for specifications using respiratory diseases as controls, while the largest are from models using non-infectious childhood disease controls, suggesting the importance of cross-disease spillovers.

The estimates for all-cause child mortality using large cities as comparators suggest an 83% relative decline in under-5 deaths by 1995 due to PAL, though this is less precisely estimated (and no longer statistically significant once we allow for clustering at higher geographic levels or when using count data models – see Section 4.3). These estimates are far larger than what we would expect if PAL only had an impact on diarrheal disease.²⁷ One explanation for this is spillover effects of PAL on diseases for which water quality is not a direct risk factor, which would imply that our estimates for diarrheal disease mortality relative to deaths from either set of control diseases would be an under-estimate of the underlying true program effect. However, this is unlikely to be the only explanation given that our models comparing diarrheal disease mortality across small and medium-sized versus large cities – which are not prone to cross-disease spillovers – yield similar estimates to our control disease approaches. A second explanation is that we under-estimate the reduction in diarrheal disease mortality due to improvements in the quality of cause of death ascertainment, which should not affect estimates for all-child child mortality. However, the magnitude of this improvement (see *Figure A1*) suggests this explanation may not fully account for the large effects noted here, either. A third possibility is that there are unobserved differences in pre-existing disease trends between larger and smaller cities, which may not be captured by a visual inspection of pre-trends. Given how much larger the estimated effect sizes are for all diseases relative to diarrheal diseases, this explanation may be most likely.

C. Extensions and Robustness

The substantive findings in *Table 3* are robust to using a quartic root transformation of the dependent variable (*Table A1*). They are also robust to modelling death counts instead of death rates using a negative binomial model, with the notable

²⁶ In an earlier working paper version of this paper, we show that diarrheal disease mortality declined in each of the age groups, 0-1 month, 1-12 months and 1-4 years (Bhalotra et al., 2018)

²⁷ Diarrheal diseases accounted for ~20% of all child deaths nationwide in the pre-PAL period. Assuming accurate cause of death registration, had the impact of PAL been restricted to diarrheal disease we would expect a 9.4%-13.4% relative decline in overall child mortality (i.e., 20% of the 47-67% relative decline in diarrheal deaths due to PAL estimated from our difference-in-difference models).

exception of models examining all-cause mortality, for which estimates in count data models imply smaller, and no longer statistically significant, effect sizes (*Table A2*). Statistical inference for models using control diseases remain robust when clustering at a higher geographic level; for models using the control city approach, estimates of both the level and trend breaks are no longer statistically significant once we cluster at the state level (*Table A3*). We also examine if migration was correlated with program impacts but we find no evidence of differential migration (*Table A4*).

IV. Under What Circumstances is Municipal Water Disinfection Effective?

The previous section demonstrates large average treatment effects of PAL. Given mixed experiences with water system chlorination/disinfection in many lower-income countries, which contrasts with the more systematically positive historical experience of wealthy countries, we next investigate the circumstances under which chlorination can be most effective. We focus on heterogeneity by coverage of municipal water pipe infrastructure, which delivers drinking water to households (and, importantly, *chlorinated* drinking water under PAL). We would expect to see larger treatment effects in areas with more extensive piped water coverage. We also examine complementarities with sanitation infrastructure, which in other work has been shown to complement water purification efforts (Alsan and Goldin, 2019, Duflo, et al., 2015).

A. Empirical Strategy

To examine heterogeneous program effects, we interact pre-intervention measures of municipal-level infrastructure from the 1990 population census (specifically, piped water coverage, sewage infrastructure coverage, and a proxy for the age of the piped water system) with each of the terms in the non-parametric and parametric difference-in-difference model using control diseases (Equations 1 and 3). Given that the main difference-in-differences results are substantively similar regardless of using respiratory or non-infectious diseases as the comparison, we estimate heterogeneous effects of PAL by baseline infrastructure using both sets of control diseases to improve power. We allow pre-existing trends, PAL program impacts, and heterogeneity in infrastructure characteristics to vary separately for each control disease.

Municipalities with higher pre-PAL population coverage of piped water and sewage infrastructure tended to be wealthier and have lower rates of diarrheal disease mortality prior to the program. This may bias interactions of the PAL exposure variable with infrastructure towards zero if, for instance, water disinfection efforts yielded higher returns in areas where disease prevalence was greater, for example.²⁸ It is therefore important to include controls for municipality economic status. We include interactions with municipality-specific average years of completed schooling, indigenous population share, and the natural logarithm of average household earnings.²⁹

B. Results

Figure 6 reports marginal effects from variants of Equation 1 examining heterogeneity in PAL program effects. Specifically, we plot estimates at the highest and lowest quartiles of piped water and sanitation infrastructure coverage. Coefficient estimates from parametric models are provided in *Table A5*. Overall, we find no statistically significant relationship between the extent of piped water coverage and the magnitude of mortality decline under PAL, with the estimated marginal effects of PAL being similar for municipalities in the lowest and highest deciles of piped water coverage (Panel A, *Figure 6*). By contrast, we find striking evidence of complementarity with sewage infrastructure (Panel B, *Figure 4*), particularly in the early years. The estimated marginal effect of PAL on diarrheal disease mortality by 1995 at the highest quartile of sewage coverage implies a 90% relative decline in diarrheal disease mortality, which is similar to historical experience. By contrast, the corresponding estimate for the lowest quartile of sewage infrastructure is a 50% relative decline over the same time period.

C. Examining the Role of Infrastructure Degradation

Our finding that there is no discernible complementarity between pipe network coverage and PAL disinfection impacts is striking because chlorinated water could only be widely delivered through water pipes. One possible explanation for this is that more

²⁸ Consistent with this possibility, in an earlier working paper version (Bhalotra et al 2018), we show that PAL induced convergence in diarrheal mortality rates between areas with high and low pre-program diarrheal disease mortality.

²⁹ Another potential confounder could be political targeting and capacity. For example, the relationship between municipality governance and national policymakers could influence the allocation of transfers and infrastructure development, a phenomenon that has been well-studied in Mexico (Diaz-Cayeros et al., 2016). However, Fried and Venkataramani (2016) find no relationship between PAL treatment effects and the pre-PAL municipality governor's political party.

extensive municipal water systems are generally older (Office of Economic Cooperation and Development (OECD), 2006, Oswald Spring, 2011, Vasquez et al., 2009). Water system age is, in turn, also correlated with the degree of infrastructure degradation in many countries (Larsen et al., 2016, Lee and Schwab, 2005, Moe and Rheingans, 2006). Consequently, age-related pipe breaks, intermittent water pressure, and resulting recontamination from the surrounding soil may have undermined the full potential health benefits of PAL (Ercumen et al., 2014, Kumpel and Nelson, 2013, Mazari-Hiriart, et al., 2005).

Although the data required to truly test this hypothesis is not generally available, two pieces of evidence support the contention that infrastructure degradation may explain the lack of heterogeneity in PAL impacts by piped water coverage. First, a large engineering literature on Mexico suggests that older municipal water systems are in fact more prone to infrastructure failure and fecal contamination (Adler, 2015, Lee and Schwab, 2005, Mazari-Hiriart, et al., 2005, World Bank, 1994).³⁰ Second, although systematic data on infrastructure degradation, water system age, and water pressure (including pressure fluctuations) are not generally available for any country in the world,³¹ we find suggestive evidence that areas with more expansive piped water coverage in 1990 do in fact tend to have older pipe infrastructure. Specifically, we construct a proxy measure for water system age by calculating the ratio of piped water coverage in the 1960 census to coverage in the 1990 census. Assuming that pre-existing piped water systems were not updated (a reasonable assumption in our context (Secretaria de Medio Ambiente y Recursos Naturales, 2013)), higher ratios imply older water systems. Consistent with this view, *Figure 7* shows that this measure of pre-existing

³⁰ For example, a recent government report estimates that 30-50% of drinking water nationwide may be lost due to pipe age, poor pressure control, and degraded materials (Secretaria de Medio Ambiente y Recursos Naturales, 2013). This illustrates how infrastructure-driven deficits in the completeness of disinfection and increases in recontamination risk are widespread.

³¹ A notable exception is the study by Ashraf et al (2017), which utilizes rich data on system outages for one district in Lusaka, Zambia. For our purposes, the best available database for worldwide municipal water system characteristics is maintained by the International Benchmarking Network for Water and Sanitation Utilities (IB-NET, <http://www.ib-net.org/>). This database includes information on total water system output and usage, infrastructure quality (including pipe breaks), and tariffs for a number of local water systems. We searched this database for information on Mexican utilities. However, data were only available for a small number of utilities and only for 2005 and onwards. Moreover, these utilities were predominantly in major cities, where PAL was less active. We use these data in some suggestive analyses (see below), but their incompleteness precludes more definitive investigation.

piped water coverage was positively associated with pipe breaks per kilometer – a measure of water system degradation – in 2005 among a sample of 16 municipal systems (the most complete data on pipe break density available).³² Assessing the relationship between our measure of water system age and piped water coverage in 1990, *Figure 8* then shows that areas with older water systems did in fact have greater piped water coverage in 1990.³³

Another potential explanation for the lack of expected heterogeneity by piped water coverage in PAL program impacts could be the impact of other program activities, e.g., provision of point-of-use chloride tables. However, as we discuss in Section 2, these program efforts were quite small relative to disinfection, and thus unlikely to explain these findings.

V. Cost Effectiveness

We have shown that PAL was associated with large reductions in diarrheal disease mortality, but a natural policy question is if this was worth the expense relative to alternative uses of public funds. To calculate the cost per child death averted and the cost per life year saved across Mexico as a whole, we use the estimates from *Table 3* for under-5 mortality rates (for respiratory controls, providing a conservative estimate). We focus on the period 1991-1995 given that we know the total cost of the program over this time frame: PAL cost approximately U.S. \$1 billion over this period (or \$1.86 billion in 2015 dollars).³⁴ Of course, the health benefits of PAL presumably continued accruing after 1995. Depending on the costs of ongoing maintenance and operation of chlorination facilities, our results may thus understate the cost-effectiveness of the program.³⁵

³² The correlation coefficient between the two variables is 0.44, $p=0.12$.

³³ We directly examining treatment effect heterogeneity by estimating versions of Equation (1) including interactions with the logged number of pipe breaks per km in 2005 for the 16 municipalities for which we have data. While the confidence intervals are large, the estimates do suggest that areas below the median of pipe breaks experienced larger PAL-led declines in diarrheal disease mortality as compared to areas above the median of this measure (*Figure A3*).

³⁴ Personal communication with Dr. Jaime Sepulveda, Mexico's Vice-Minister of Health from 1991 to 1994.

³⁵ Another reason our estimates may be conservative is that we do not account for additional program benefits beyond life years saved. For example, other work demonstrates robust short run benefits of water infrastructure on short-run financial transactions (Ashraf et al, 2017) and time allocations and long-run benefits of infant exposure to PAL on cognitive outcomes and schooling for girls (Bhalotra and Venkataramani, 2015).

For each year between 1991 and 1995, we calculate the differential reduction in diarrheal disease mortality rates using the estimates for the level shift ($1(Diarrhea) \times 1(Post)$) and trend break ($1(Diarrhea) \times 1(Post) \times Year$) from each of our difference in differences models. We multiply these estimated rate changes by the average number of live births in small and medium-sized towns (treated areas) in each year prior to PAL to recover the implied number of averted deaths. We then sum the implied number of averted deaths across the years 1991-1995. The results of this exercise are presented in *Table 4*. Across specifications, we estimate that PAL averted between 9,140 and 16,437 child diarrheal deaths between 1991-1995 at a cost per death averted of \$113,782-\$203,501. Assuming that individuals surviving childhood would have lived to the age of 60 (conservative given that the average life expectancy at the time of the policy was 71 years), the cost per life year saved was \$1,896-\$3,392.³⁶

To estimate cost-effectiveness given the presence of optimal complementary infrastructure, we use the estimates from Section 5.2 to assess the expected decline in deaths for the counterfactual scenario in which all municipalities are assigned the top quartile value of sewage coverage in our sample. The 90% decline in diarrheal disease mortality by 1995 predicted for this scenario implies over 26,000 averted child diarrheal deaths between 1991-1995, yielding a cost per life year saved of \$1,170. However, this estimate does not account for the cost of building complementary infrastructure. To allow for this, we use data from a World Health Organization report on the average per capita cost of extending and improving water and sanitation coverage and the cost of maintaining good quality infrastructure access in a sample of 94 low- and middle-income countries (Hutton, 2012) between 2005-2014 (estimates specifically for Mexico for the time period of our study are not available). These estimates imply an outlay of \$5.1 billion over a 5-year period. Taking these costs into account – and assuming that they translate immediately into improved infrastructure – implies a cost of \$4,388 per life year

³⁶ Models comparing child mortality from all causes across small-and medium-sized versus large cities yield an estimate of over 120,000 child deaths averted by 1995, at a cost per life year saved of \$215. As discussed in Section 4.2, these estimates are likely far larger than what we would expect from fully accounting for the Mills-Reinke effect and from improvements in cause of death attribution – and consequently may reflect instead unobserved pre-existing differential epidemiologic trends between larger vs. small- and medium-sized cities.

saved.³⁷ This estimate is still relatively low, and is also likely an upper bound given that it does not account for the potential direct consequences of these investments on morbidity and mortality as well as cross-disease spillover effects.

VI. Discussion and Conclusion

Mexico's *Programa Agua Limpia* provides an unusual opportunity to study the effects of a contemporary, state-funded municipal water disinfection effort – and to examine within-country variation in program effects to study under what circumstances disinfection can be differentially effective in improving health. The importance of this case is underscored by the fact that, at the time of the program (the early 1990s), childhood diarrheal disease mortality rates in Mexico were higher than in the average low-income country today. We find that the program reduced childhood diarrheal disease mortality rates by 47-67% nationwide and was cost-effective, with a cost per life-year saved of \$1,896-\$3,392. Despite the fact that piped water infrastructure was required to distribute water disinfected with chlorine, we also find that PAL's reductions in diarrheal disease mortality rates were no larger in municipalities with greater baseline piped water coverage rates. In exploring this result, we find suggestive evidence that poor quality water infrastructure (e.g., breaks in piped water infrastructure) - which is more prevalent in older systems with greater population coverage – may have dampened the impacts of disinfection. In contrast, we find significant complementarities between sewage infrastructure and the PAL program, consistent with prior work.

More generally, these results imply that maximizing the return to large-scale water quality interventions in low- and middle-income countries may require concurrent infrastructure investments, including repairing degraded water pipes and expanding

³⁷ Specifically, the World Health Organization estimates the annual expenditures required over the period 2005-2014 across 94 low- and middle-income countries to meet access and quality targets set by the United Nation Millennium Development Goals. The estimates include the cost of extending piped water and sanitation coverage to those not currently receiving them, accounting for population growth, as well as the cost of updating and maintaining existing facilities. See Hutton (2012) for further details. The resulting annual outlay estimated is \$12 per capita. For our cost effectiveness estimates, we assume that these costs are borne annually over the five post-PAL years that we study (i.e., \$60/capita over five years multiplied by 85.4 million, the 1990 population of Mexico). This yields an estimate of upgrade costs of \$5.1 billion dollars over 1991-1995. Added to the costs of PAL, the estimated total outlay is \$6.96 billion. Using our estimates for PAL treatment effects for municipalities in the top infrastructure decile of 90%, we estimate 26,347 deaths averted as a result. This yields a cost-effectiveness estimate of \$4,388 per life year saved.

sewage infrastructure. This is underscored by the stark positive correlation between direct measures of infrastructure quality and diarrheal disease rates among low- and middle-income countries around the world (*Figure 9*). Although infrastructure improvement and expansion is expensive, our back-of-the-envelope calculations for Mexico suggest that even when including them, water disinfection may still be cost effective, with a cost per life-year saved of approximately \$4,400 – and with many more lives saved in total. Overall, we emphasize the complexity of contemporary urban water and sanitation challenges in developing countries and the need for more research on how to improve the safety of municipal drinking water, even with otherwise proven technologies.

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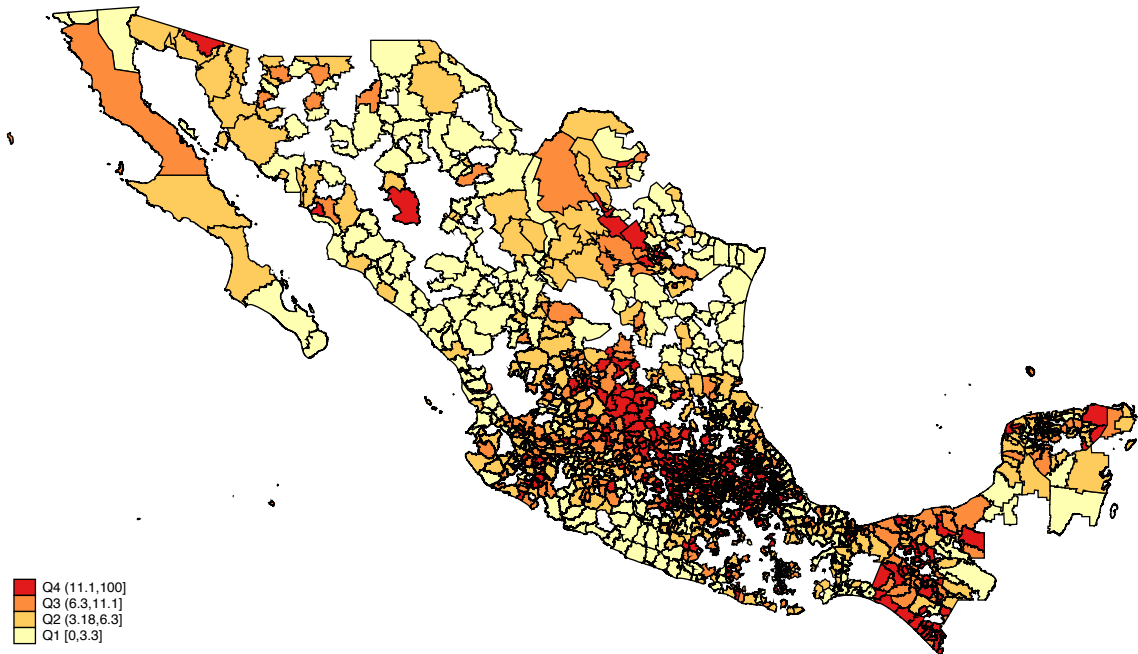
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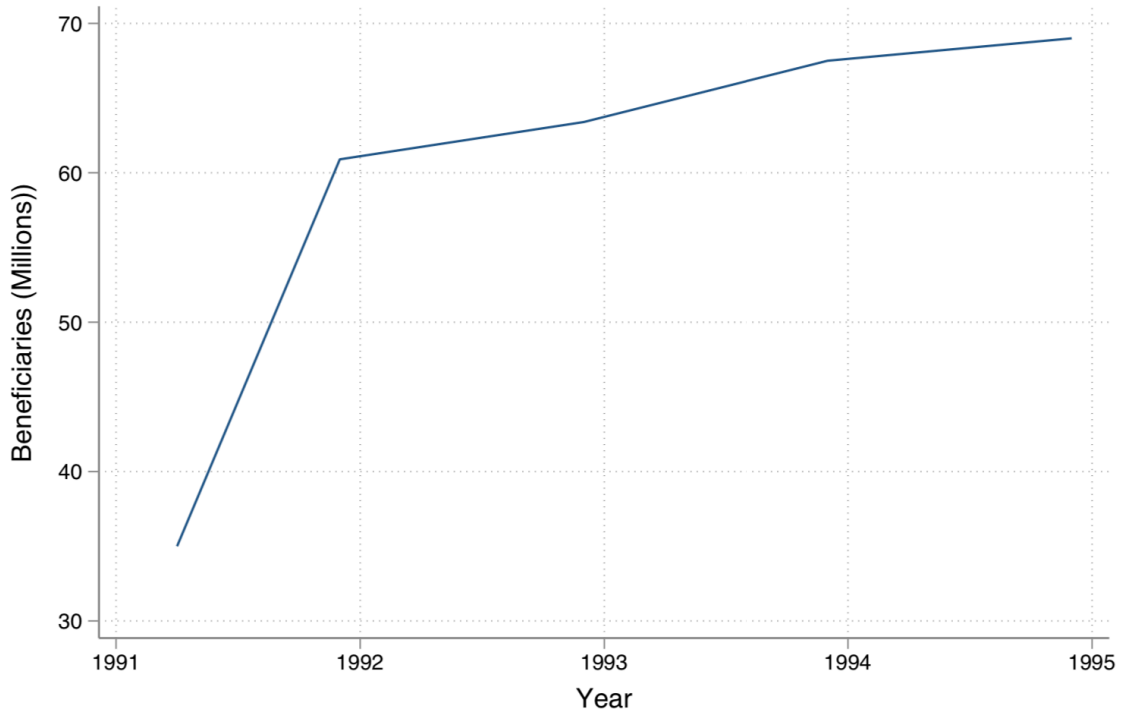
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Figure 1: Pre-Intervention Under-5 Diarrheal Mortality Rates Across Urban Mexican Municipalities (1985-1990 average)



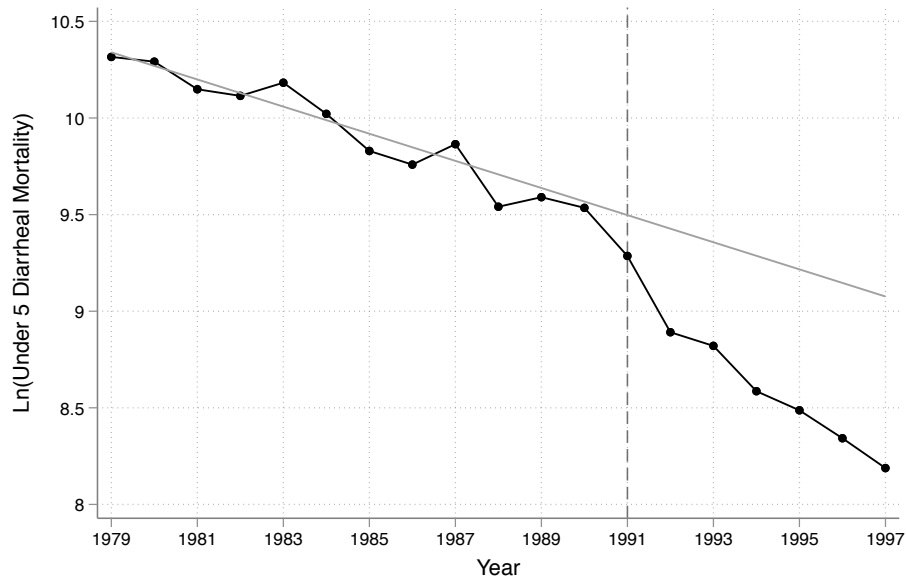
Notes: Heat map plots quartiles of diarrheal mortality rates per 1,000 live births over the period 1985-1990 for children under the age of 5 by municipality. Darker colors reflect higher average pre-intervention diarrheal mortality rates. Fully rural municipalities are excluded from the map (as they are in our main analyses). Data to construct map were obtained from the Mexico Ministry of Health, Vital Statistics.

Figure 2: Chlorination Coverage under Programa Agua Limpia



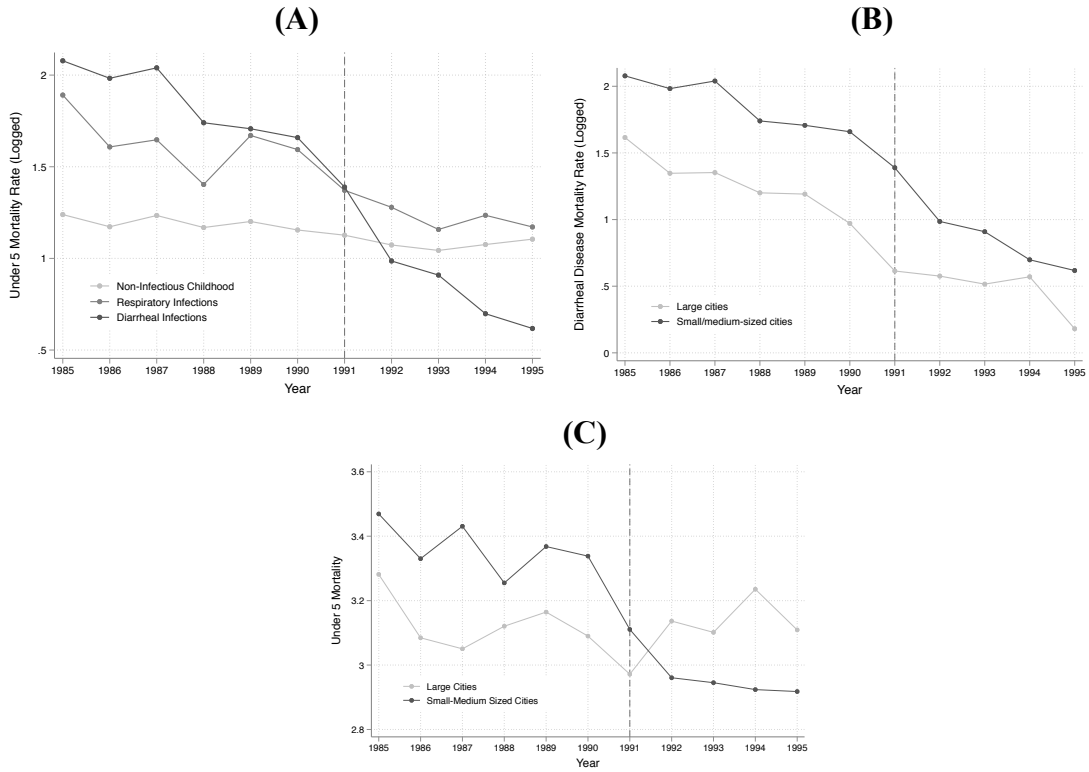
Notes: Figure plots the number of beneficiaries (in millions) covered by chlorination services between April 1991 and December 1995. The sharp uptick in coverage between April 1991 and December 1991 was coincident with the introduction of Programa Agua Limpia (PAL). These program efforts correspond to an increase in the share of the urban population covered by disinfected water from 58% to over 90% of the urban population covered by disinfection (or from 47% to 72% for the full population). Data were obtained from CONAGUA (1994).

Figure 3: National Trends in Under-5 Mortality Diarrheal Deaths, 1979-1997



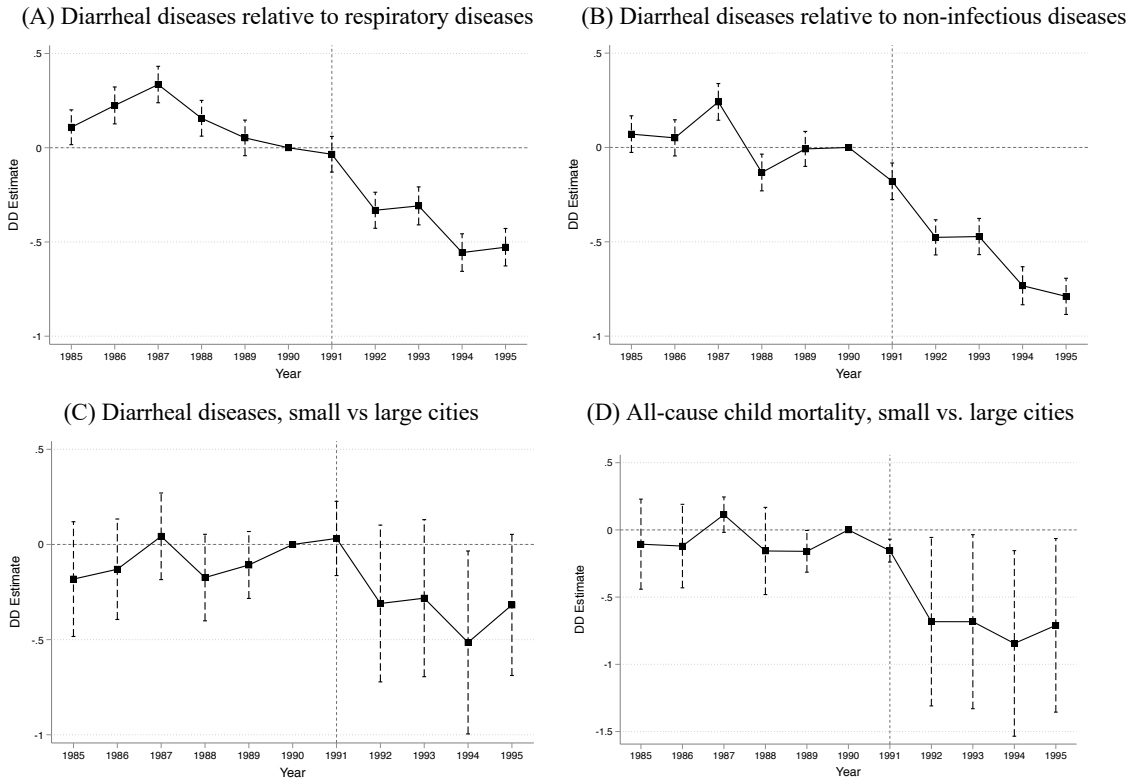
Notes: Figure plots the average trends in mortality rates for diarrheal diseases for non-rural municipalities between 1979-1997. Visually, we note a trend break in the diarrheal mortality series in 1991, coincident with the start of PAL. The grey line is the linear prediction of the pre-PAL trend between 1979 and 1984 – notably, actual diarrheal mortality followed this prediction till 1991, when PAL was implemented. This break was confirmed econometrically using the Quandt Likelihood Ratio test ($F = 7.82$, $p = 0.013$) assessing for the presence of breaks between 1985-1995). Breaks for the control diseases of respiratory infections and non-infectious childhood conditions were either not statistically significant or not timed with policy implementation (see *Figure 4* and *Figure A2*). Data source: Ministry of Health, Vital Statistics.

Figure 4: National Trends in Under-5 Mortality Rates by Treated and Control Locations or Conditions, 1985-1995



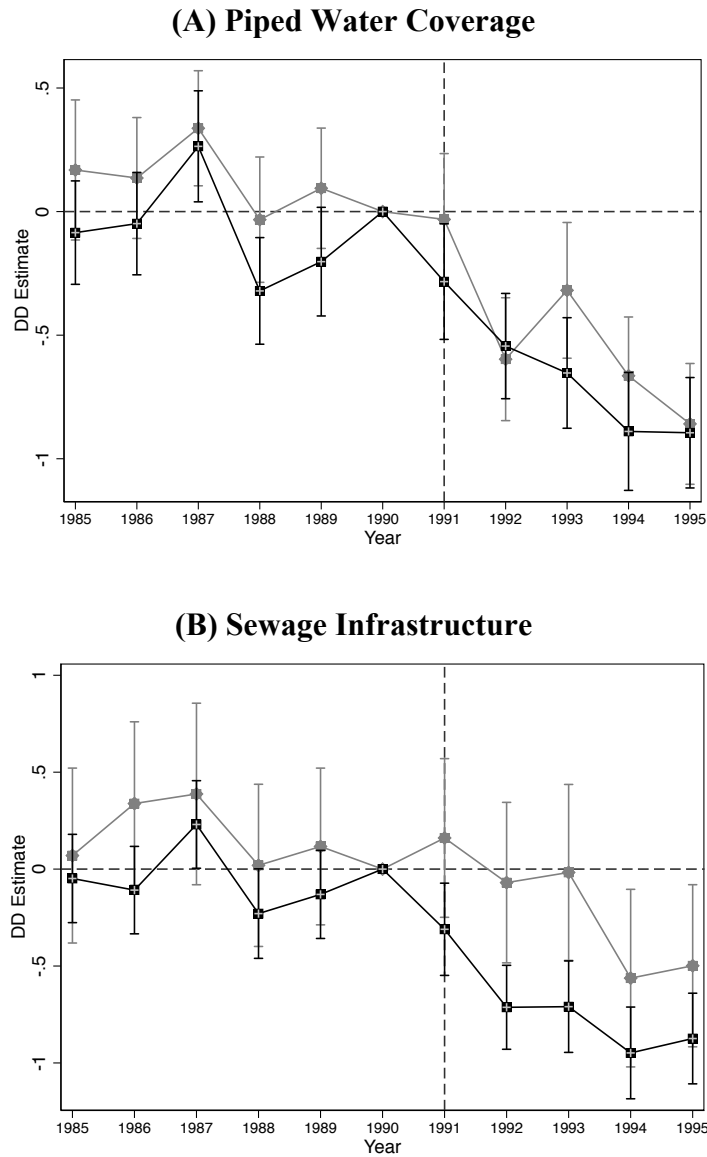
Notes: Figure plots the average trends in under 5 mortality rates (per 1,000 live births) for diarrheal diseases alongside respiratory and non-infectious control diseases in municipalities containing small- and medium-sized cities (panel A), diarrheal diseases in municipalities containing small- and medium-sized cities versus large (untreated) cities (panel B), and all-cause under-5 mortality across municipalities containing small- and medium-sized cities versus large cities (Panel C).

Figure 5: Event Study Plots, Under 5 Diarrheal and All-Cause Child Mortality Rates Relative to Controls



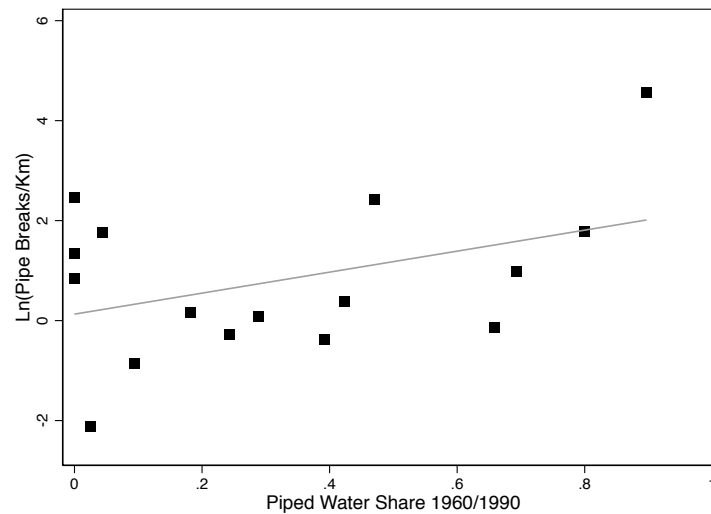
Notes: Panels plot event study estimates of δ_t (Panels A and B) and at (Panels C and D) from Equations (1) and (2), respectively. Panels A and B compare diarrheal disease mortality rates to non-infectious childhood diseases and respiratory infection controls, respectively. Panels C and D compares under-5 diarrheal disease mortality rates and all-cause under-5 mortality rates, respectively, in small-and medium-sized cities versus large cities, the latter of which already had chlorinated water supplies well in advance of PAL. Coefficients are denoted by the dots and the vertical line and whiskers denote the 95% confidence interval of the estimates. The dashed lines are predicted trend lines calculated (separately) using pre-1991 coefficients and 1991-1995 coefficients. These trend lines reveal a distinct trend break starting in 1991.

Figure 6: Heterogenous Treatment Effects by Pre-PAL Levels of Infrastructure



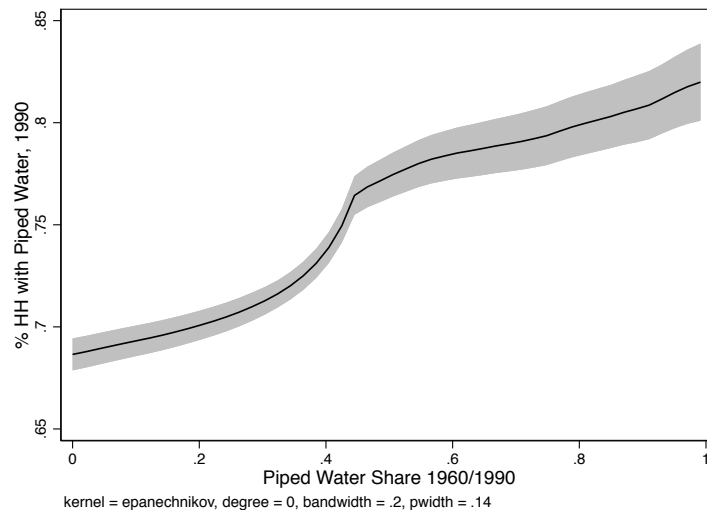
Notes: Panels plot event study estimates of PAL impacts on under-5 diarrheal disease mortality rates from versions of Equation (1) that include the full set of interactions with baseline municipality infrastructure coverage measures (share of population with piped water and share of population with access to sewage infrastructure) and socioeconomic covariates (average years of schooling among adults, mean household income (logged), and share of indigenous population). The models use both respiratory and non-infectious childhood diseases as control conditions in order to maximize statistical power. Panel A plots marginal effect estimates at the bottom (grey) and top (black) quartiles of piped water infrastructure and Panel B plots marginal effect estimates at the bottom (grey) and top (black) quartiles of sewage infrastructure. Differences in treatment effects are only statistically significant for sewage infrastructure (see *Table A5*).

Figure 7: Relationship Between Water System Age Proxy and Pipe Main Breaks



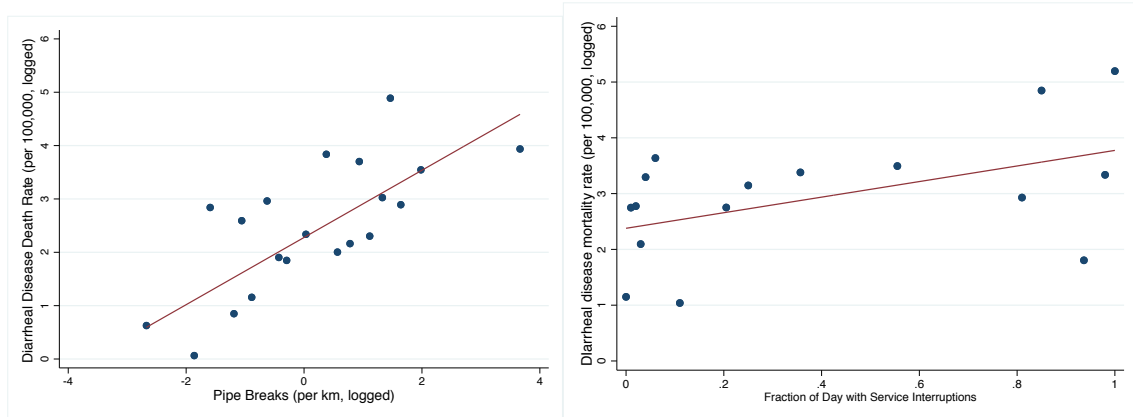
Notes: Scatter plot of logged pipe breaks per kilometer in 16 municipal water systems (distributed across 11 states) in 2005 against the ratio of households with access to piped water in (1960/1990), which we use as a proxy for system age. Data on pipe breaks were obtained from the International Benchmarking Network for Water and Sanitation Utilities (IB-NET, <http://www.ib-net.org/>). The best-fit line reveals a positive correlation between these variables ($\rho = 0.41$, $p=0.12$).

Figure 8: Water System Age and 1990 Population Coverage



Notes: LOWESS smoothed plots (with 95% CI) of municipality-specific ratios of household piped water coverage between 1960 and 1990 (our proxy for water system age) and piped water coverage in 1990. The positive slope of the plot suggests that systems with higher population coverage pre-PAL may also have been older and, therefore, dilapidated. *Figure 7* directly makes the case for the latter point, linking the 1960/1990 access ratio to pipe breaks in 2005.

Figure 9: Infrastructure Quality and Diarrheal Disease Mortality Rates in Low- and Middle-Income Countries



Notes: Ventile plots correlating measures of infrastructure quality in 2015 (water pipe breaks per km; fraction of an average day with water service interruptions) with (all age) diarrheal disease mortality rates (per 100,000 population) in the same year for a sample of 42 low- and middle-income countries for which all measures were available. Infrastructure data were obtained from the International Benchmarking Network for Water and Sanitation (IB-NET, <http://www.ib-net.org>). Diarrheal disease mortality rates were obtained from the Institute of Health Metrics and Evaluation, via the Our World in Data Website (Diarrheal diseases death rates (<https://ourworldindata.org/diarrheal-diseases>)).

Table 1: Descriptives for Vital Statistics

	Pre-PAL (1985-1990)		Post-PAL (1991-1995)		Full Sample	
	Mean	S.D.	Mean	S.D.	Mean	S.D.
<i>Municipalities with Small- and Medium-Sized Cities (Targeted by PAL chlorination)</i>						
Diarrheal Mortality Rate (Treated)	6.5	7.2	2.6	3.3	4.7	6.0
Respiratory Mortality Rate (Control)	5.2	7.6	3.5	4.8	4.4	6.5
Non-Infectious Childhood Disease Mortality Rate (Control)	3.3	3.9	3.0	2.7	3.1	3.4
All-Cause Mortality Rate	28.4	26.6	19.1	14.4	24.1	22.3
<i>Municipalities with Large-Cities (Not Targeted by PAL chlorination)</i>						
Diarrheal Mortality Rate	3.7	3.9	1.7	1.0	2.8	3.2
All-Cause Mortality Rate	23.0	24.7	22.4	15.4	22.8	21.2

Notes: Descriptive statistics for 1,413 municipalities containing small- and medium-sized cities, which were treated by PAL chlorination efforts, and 16 municipalities containing large cities, which were not targeted by PAL chlorination. We focus specifically on quantities relevant to our key difference-in-difference comparisons. Thus, among the first group, we present means for infectious diarrheal disease mortality rates (ICD9 codes 001-009), for which clean water is a risk factor, and mortality rates from respiratory (upper and lower respiratory infections - ICD9 codes 60-466, 480-487) and non-infectious childhood diseases (the sum of death rates from congenital anomalies (ICD9 codes 740-759) and perinatal causes (low birth weight, birth trauma, congenital infections, neonatal jaundice, etc: ICD9 codes 764-779), for which clean water is not a direct risk factor. We also present data for all-cause under-5 mortality rates. For municipalities with large cities, we present diarrheal disease and all-cause mortality rates. All statistics are weighted by the number of live births. For the pre-PAL period, statistics for small- and medium-sized (large) cities are based on observations from 8,478 (96) municipality-years. For the Post-PAL period, statistics for small- and medium-sized (large) cities are based on observations from 7,063 (80) municipality-years.

Table 2: Descriptives for Infrastructure and Socioeconomic Characteristics

Municipality Characteristics, 1990 Census	Mean	S.D.
% HH with Access to Piped Water	69.5	21.5
% HH with Access to Sewage	36.6	22.5
% Adults Completing Secondary Schooling	5.4	1.58
Average Per Capita Earnings (Pesos, 1000)	27,106	12,247

Notes: Characteristics for municipalities containing small-and medium-sized cities from 1990 (taken as a pre-PAL baseline) were computed using the IPUMS 1990 Census 10% micro data (Minnesota Population Center, 2015), aggregated (using appropriate weights) to the municipality level. Data were available for 1,282 of the 1,413 municipalities containing small-and medium-sized cities. The piped water variable reflects the percentage of households in a municipality with access to piped water either at their home or via a public tap. The sewage variable reflects the percentage of households with on premises access to a sewage system. The earnings variable is the mean total annual household earned income in 1990 pesos (the large numbers reflect significant currency devaluation over the 1980s-early 1990s).

Table 3: Differences-in-Differences Estimates
Dependent Variable: Under-5 Mortality Rate

Outcome	Diarrheal Diseases	Diarrheal Diseases		Diarrheal Diseases	All Diseases
<i>Control Group</i>	<i>Respiratory Diseases</i>	<i>Non-Infectious Diseases</i>		<i>Large Cities</i>	<i>Large Cities</i>
<i>1(Diarrhea)*1(Post)</i>	-0.132 (0.0418)	-0.182 (0.0425)	<i>1(Small)*1(Post)</i>	0.00262 (0.0361)	-0.303 (0.153)
<i>1(Diarrhea)*1(Post)*Year</i>	-0.0857 (0.0144)	-0.122 (0.0143)	<i>1(Small)*1(Post)*Year</i>	-0.112 (0.0381)	-0.131 (0.0790)
<i>1(Diarrhea)*Year</i>	-0.0354 (0.00839)	-0.0258 (0.00872)	<i>1(Small)*Year</i>	0.0217 (0.0300)	0.00408 (0.0377)
<i>1(Post)</i>	-0.146 (0.0299)	-0.0966 (0.0296)	<i>1(Post)</i>	-0.196 (0.130)	-0.0314 (0.150)
<i>1(Post)*Year</i>	-0.0237 (0.0108)	0.0124 (0.00993)	<i>1(Post)*Year</i>	0.00262 (0.0361)	0.0543 (0.0781)
<i>Year</i>	-0.00964 (0.00675)	-0.0192 (0.00647)	<i>Year</i>	-0.0668 (0.0289)	0.0279 (0.0368)
<i>1(Diarrhea)</i>	0.218 (0.0374)	0.962 (0.0374)			
N	31,082	31,082	N	15,717	15,717
R-squared	0.471	0.437	R-squared	0.485	0.563
<i>% Decline by 1995 Due to PAL</i>	-47%	-67%	<i>% Decline by 1995 Due to PAL</i>	-45%	-83%

Notes: Estimates of Equations (3) and (4). Robust standard errors, correcting for clustering at the municipality level, are provided in parenthesis. All models include municipality fixed effects. Each column represents a separate regression, with the column headers denoting the under-5 mortality rate outcome of interest and comparison groups for difference-in-differences model. All dependent variables, which are mortality rates for children under age 5, are transformed using the inverse hyperbolic sine function. In the first two columns, which present estimates for models using control diseases, $1(Diarrhea) = 1$ denotes diarrheal disease mortality rates while $1(Diarrhea) = 0$ equals mortality from the control diseases. In the last two columns, which present estimates for models using municipalities containing large cities are controls, $1(Small) = 1$ denotes observations from municipalities containing or located within small and medium-sized cities (<500,000 population) versus large cities (>500,000) that were not treated by PAL chlorination efforts. In both panels, $1(Post)$ indicates if the year of observation is 1991 or thereafter. Across all models, the first two terms represent estimates of the level and trend breaks, respectively, and are used to generate treatment effect estimates (*% Decline by 1995 Due to PAL*). The terms $1(Diarrhea)*Year$ and $1(Small)*Year$ and capture differential pre-existing trends in the outcome compared to control diseases or control areas, respectively. The sample size reflects the number of municipality-years.

Table 4: Cost Effectiveness

	Under-5 Diarrheal Deaths Averted	Program Cost (2015 USD)	Cost/Death Averted (2015 USD)	Cost/Life Year Saved (2015 USD)
Respiratory Disease Controls	12426	1.86 billion	149686	2495
Non-Infectious Disease Controls	16347	1.86 billion	113782	1896
Large City Controls	9140	1.86 billion	203501	3392

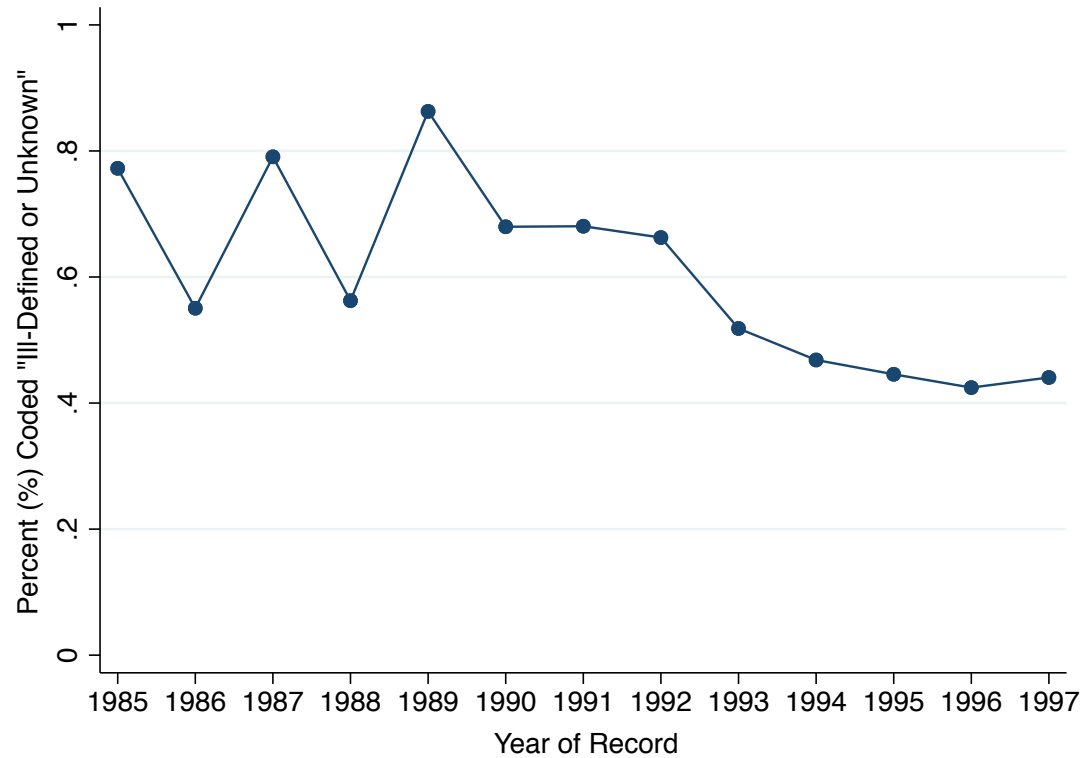
Notes: ~2.1 million live births occurred each year between 1985-1990 in small and medium-sized municipalities. We use this figure, the pre-intervention average baseline mortality rate from diarrheal disease, and estimates from the difference-in-differences models examining diarrheal disease mortality as an outcome from **Table 3** to calculate the number of under-5 deaths averted each year for each outcome between 1991-1995. The column "Under-5 Diarrheal Deaths Averted" displays these estimates. The total program cost – 1.86 billion USD - over this period comes from administrative sources. The cost per death averted is simply the total cost/number of deaths averted. To calculate Cost/Life Year Saved, we assume each under-5 child whose death was averted would (conservatively) live to the age of 60 (and so we multiply the numbers in the penultimate column by 60 to get the numbers in the last column). All costs are expressed in 2015 USD.

ONLINE APPENDIX

Urban Water Disinfection and Mortality Decline in Lower-Income Countries

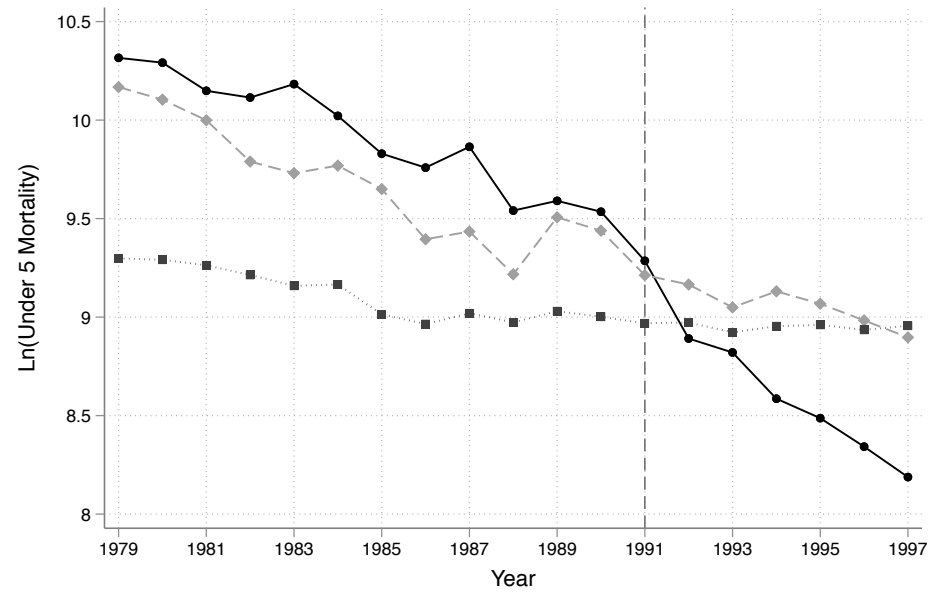
Sonia R. Bhalotra, Alberto Diaz-Cayeros, Grant Miller, Alfonso Miranda, and Atheendar S. Venkataramani^{||}

Figure A1 – Quality of Death Registration Records



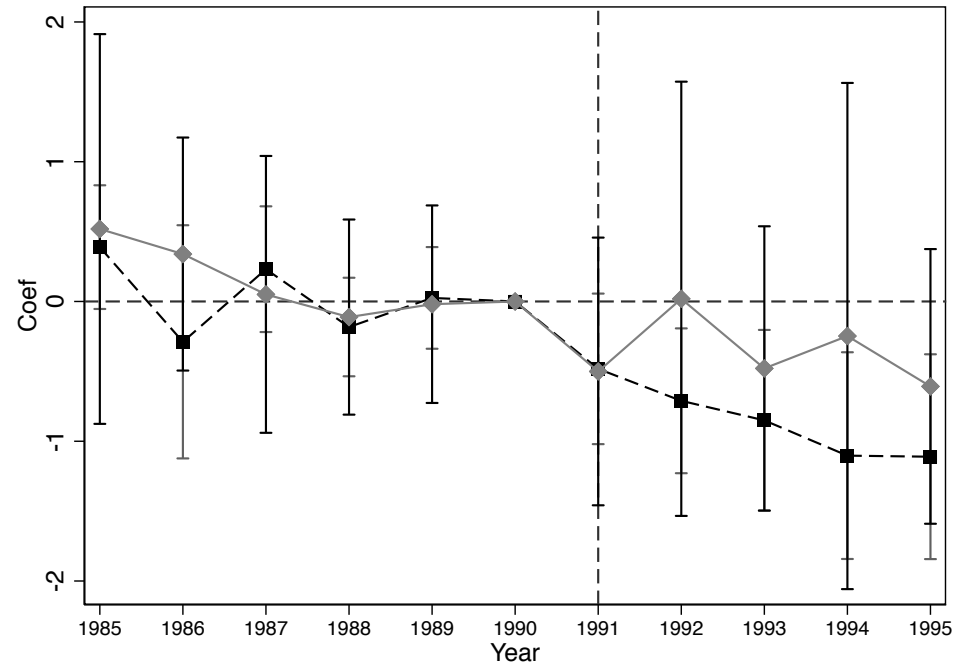
Notes: We conducted two separate analyses to assess the quality of Mexico’s vital registration records during our study period. First, we examined the number of death records for which the cause of death was missing or not coded, finding that 0% of death records had missing causes. Second, we assessed the prevalence of causes of death coded as “ill-defined or unknown causes of mortality” over time, which we plot in the above Figure. While the share does decline slightly after 1990, the prevalence of unknown cause codes is less than 1% through the period. We note that these analyses do not address the potential for changes in the completeness of recording deaths from any cause and more accurate attribution and assignment of causes of death over time.

Figure A2 – National Trends in Under-5 Deaths for Control Diseases



Notes: Figure plots national trends logged under 5 deaths from diarrheal diseases (black circles), respiratory control diseases (light grey diamonds), and non-infectious diseases (dark grey squares) for the period 1979-1997 using data from the Mexican Vital Statistics registry. We plot logged number of deaths instead of death rates here because municipality level data on the number of births each year are not available prior to 1985. As noted in the main text, we formally tested for structural breaks for each disease between 1985-1995 (Quandt 1960), remaining agnostic about the exact break point in the time series. Specifically, we calculate the F -statistic on different user-specific break points in the window, with the largest F -statistic across tests of different time points is used to identify the break point. We ran these tests after detrending the time series using the estimated linear time trend over 1979-1985, after which the null hypothesis of a unit root in the time series prior to any structural breaks was rejected by an augmented Dickey-Fuller test (e.g., see Hansen 2001). We found a statistically significant trend break for diarrheal diseases in 1991, which is timed exactly with PAL ($F = 7.82$, $p = 0.013$). We do not find evidence of any trend breaks timed with PAL for either respiratory diseases (break year 1987, $F = 2.76$, $p=0.12$) or non-infectious childhood diseases (break year 1989, $F = 6.92$, $p=0.018$).

Figure A3 – Heterogeneous Treatment Effects by Pipe Breaks (Measured in 2005)



Notes: Panels plot event study estimates of PAL impacts on under-5 diarrheal disease mortality rates from versions of Equation (1) that include interactions with a direct measure of piped water infrastructure quality, pipe main breaks per kilometer of piped water infrastructure. These data, described in the notes to **Figure 7**, were obtained from the International Benchmarking Network for Water and Sanitation Utilities (IB-NET, <http://www.ib-net.org/>) and are available for only 16 municipal water systems (compared to 1,429 in our main analyses) and only for the year 2005. Consequently we treat these estimates as suggestive. The figure plots marginal effect estimates for municipalities below (black) and above (grey) the median of the pipe breaks measure. Treatment effect estimates are larger in magnitude where pipe breaks are lower. However, these differences are not statistically significant given the small sample size.

Table A1 – Difference in Difference Estimates Using Quartic Transform of Mortality Rates

Outcome	Diarrheal Diseases		Diarrheal Diseases		All Diseases
<i>Control Group</i>	<i>Respiratory Diseases</i>	<i>Non-Infectious Diseases</i>	<i>Large Cities</i>	<i>Large Cities</i>	<i>Large Cities</i>
1(Diarrhea)*1(Post)	-0.065 (0.025)	-0.088 (0.026)	1(Small)*1(Post)	-0.056 (0.086)	-0.167 (0.083)
1(Diarrhea)*1(Post)*Year	-0.045 (0.009)	-0.066 (0.009)	1(Small)*1(Post)*Year	-0.078 (0.022)	-0.067 (0.045)
N	31,082	31,082	N	15,717	15,717
R-squared	0.44	0.3	R-squared	0.45	0.54

Notes: Models are identical to those presented in *Table 3* of the main text, except here we use the quartic root transform of instead of the inverse hyperbolic sine. Robust standard errors, correcting for clustering at the municipality level, are provided in parenthesis. All models include municipality and year fixed effects. Importantly, the coefficient estimates cannot be interpreted in the same way as a log transform or inverse hyperbolic sine. Given difficulties in getting generalized linear model versions of these regression to converge, we were unable to calculate marginal effects for these coefficients. Nevertheless, we note that the substantive findings remain similar to those presented in the main text.

Table A2 – Difference in Difference Estimates Using Death Counts

Outcome	Diarrheal Diseases			Diarrheal Diseases	All Diseases
<i>Control Group</i>	<i>Respiratory Diseases</i>	<i>Non-Infectious Diseases</i>		<i>Large Cities</i>	<i>Large Cities</i>
1(Diarrhea)*1(Post)	-0.082 (0.026)	-0.182 (0.030)	1(Small)*1(Post)	-0.010 (0.158)	-0.072 (0.085)
1(Diarrhea)*1(Post)*Year	-0.119 (0.010)	-0.121 (0.014)	1(Small)*1(Post)*Year	-0.122 (0.059)	-0.036 (0.038)
N	31,082	31,082	N	15,717	15,717
<i>% Decline by 1995 Due to PAL</i>	<i>-59%</i>	<i>-67%</i>		<i>-50%</i>	<i>-21.6%</i>

Notes: This table is identical to *Table 3* in the main text except here we data on death counts between 1985-1995. We use a negative binomial model to model the number of deaths. We use the estimated coefficients to calculate the percent relative decline in diarrheal deaths. These estimated effects are substantively similar except for models examining deaths from all-causes, for which estimates suggest smaller (and statistically insignificant) effects. Robust standard errors, correcting for clustering at the municipality level, in parenthesis.

Table A3 – Statistical Inference After Clustering at Higher Geographic Levels

Outcome	Diarrheal		Diarrheal		All
<i>Control Group</i>	<i>Respiratory</i>	<i>Non-Infectious</i>		<i>Large Cities</i>	<i>Large Cities</i>
<i>1(Diarrhea)*1(Post)</i>	-0.132	-0.182	<i>1(Small)*1(Post)</i>	0.00262	0.0543
P-value, municipality clustering	0.002	0.000	P-value, municipality clustering	0.54	0.046
P-value, state clustering	0.004	0.000	P-value, state clustering	0.53	0.110
P-value, wild cluster bootstrap, state-level	0.002	0.005	P-value, wild cluster bootstrap, state-level	0.67	0.53
<i>1(Diarrhea)*1(Post)*Year</i>	-0.0857	-0.122	<i>1(Small)*1(Post)*Year</i>	-0.112	-0.131
P-value, municipality clustering	0.000	0.000	P-value, municipality clustering	0.003	0.096
P-value, state clustering	0.004	0.003	P-value, state clustering	0.062	0.21
P-value, wild cluster bootstrap, state-level	0.005	0.000	P-value, wild cluster bootstrap, state-level	0.19	0.53

Notes: Estimates of Equations 3 and 4 of the main text with p-values obtained from clustering at the municipality-level (as in *Table 3* of the main text) and the state-level. For state-level clustering, we additionally implement the wild cluster bootstrap-t method of Cameron, Gelbach, and Miller (*Review of Economics and Statistics* 90(3), 2008). Point estimates are the same as those presented in those presented in *Table 3* of the main text.

Table A4 – Diarrheal Mortality Decline and In-Migration

	(1)	(2)	(3)
ΔDiarrhea	0.00014 (0.000048)	0.000192 (0.000046)	0.0002 (0.000074)
N	498	498	498
R-squared	0.002	0.133	0.238
BaseDiar	0.00016 (0.00068)	0.00021 (0.00059)	0.00055 (0.00075)
N	498	498	498
R-squared	0.0001	0.106	0.201
Controls			
Municipality Char	No	Yes	Yes
State FE	No	No	Yes

Notes: To assess potential non-random migration as a function of exposure to PAL, we use data from the public use 1995 Mexican Population Census Microdata, a representative 1.5% sample which allows us to identify individuals who lived in a different municipality 5 years prior to survey (i.e., pre-PAL, which was in 1991). In prior censuses, only interstate migration is identifiable. The identity of the municipality the individual moved from is not known, though the current municipality of residence is recorded. Focus on reproductive age adults (i.e., men and women ages 18-40), we can construct the in-migration rate for each municipality (in our data we can identify nearly 500 municipalities). On average, the share of individuals living in a given municipality who migrated from elsewhere over the preceding 5-year period was 9.3%.

To assess whether in-migration responded to changes in diarrheal disease environment, we estimated models of the following form:

$$Migration_{ij} = \alpha_0 + \alpha_1 (\Delta Diarrhea_j) + X_{(i)(j)} + u_{ijt}$$

$$Migration_{ij} = \alpha_0 + \alpha_1 (BaseDiar_j) + X_{(i)(j)} + u_{ijt}$$

Here, i represents the municipality and j the state. $Migration_{ij}$ is the proportion of individuals in a given county in 1995 who lived in another county 5 years prior; $\Delta Diarrhea_j$ is the change in under-5 diarrheal disease mortality in the post versus pre-periods (defined so that positive values reflect larger declines); $BaseDiar_j$ is pre-intervention baseline diarrheal mortality rate change; and $X_{(i)(j)}$ represent municipality specific, pre-intervention controls and/or state-fixed effects.

The first regression assesses whether in-migration changed as a function of the degree of decline in diarrheal mortality. We find that areas with larger declines in diarrheal mortality had higher rates of immigration. While the estimates are precise, they are substantively small. The estimates suggest that the average drop in diarrheal mortality pre-post PAL was associated with a small 0.08% point increase in the proportion of in-migrants, which is less than 1/100th of the mean. These small estimates are robust to the inclusion of controls. The second regression leverages the insight that areas with higher pre-intervention diarrheal mortality rates gained more from PAL (see our original working paper, Bhalotra et al (2018) for further details). Here, too, we find small and, this time, imprecisely estimated coefficients. For example, at the mean of baseline diarrheal mortality rate, we would expect only a 0.16% pt increase in in-migration. We conclude that nonrandom migration is unlikely to be driving our findings.

Table A5 – Parametric Estimates of Heterogeneous Effects of PAL by Pre-Program Infrastructure

Model estimates	
1(Diarrhea)*1(Post)	0.026 (0.106)
1(Diarrhea)*1(Post)*Year	-0.0937 (0.0377)
Piped Water Covg*1(Diarrhea)*1(Post)	-0.0704 (0.221)
Piped Water Covg*1(Diarrhea)*1(Post)*Year	-0.0519 (0.0786)
Sewage Covg*1(Diarrhea)*1(Post)	-0.581 (0.272)
Sewage Covg*1(Diarrhea)*1(Post)*Year	-0.0398 (0.0954)
Municipality-Disease-Year Obs (Municipalities)	42,234 (1,280)
R-squared	0.41

Notes: Estimates from a variant of Equation 3 in the main text, where each parametric term is interacted with pre-PAL municipality infrastructure characteristics obtained from a 10% sample of the 1990 census. These estimates complement the non-parametric models examining treatment effect heterogeneity presented in *Table 6* of the main text. Models include both sets of control diseases to improve power (allowing pre-existing trends, PAL program impacts, and heterogeneity in infrastructure characteristics to vary separately for each control disease). Models also include full sets of interactions with pre-PAL municipality average (logged) household income, years of schooling completed, and share of population indigenous, also obtained from the 1990 census. We include the additional control variables in order to better identify the true interaction effect between baseline amenities and disinfection. This is because, in a cross-section, areas with better baseline amenities on average are better off, having better socioeconomic outcomes and exhibiting lower rates of diarrheal disease mortality. Failing to account for socioeconomic characteristics may lead to

falsely attenuated estimates of complementarities between infrastructure and disinfection, because the infrastructure may reflect baseline risk: that is, we would expect areas with lower baseline risk – which on average have better piped water and sewage infrastructure – to benefit less from a clean water intervention.

We only report the specific coefficients that capture main program effects ($1(Diarrhea) \times 1(Post)$ for the level break and $1(Diarrhea) \times 1(Post) \times Year$ for the trend break) and coefficients estimating the interactions between main program effects and pre-existing piped water and sewage coverage, respectively. Standard errors, clustered at the municipality level, are provided in parentheses.

Of note, the number of municipalities represented in our sample ($n = 1,280$) is smaller than the number of municipalities in our main analyses ($n = 1,429$), owing to the missing information on infrastructure characteristics for smaller municipalities who may not have been represented in a 10% random sample of the 1990 census.