Urban Water Disinfection and Mortality Decline in Developing Countries

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Non-Technical Summary

Improvements in the quality of drinking water made a significant contribution to population health in the late 19th and early 20th century in today's richer countries. Children were especially vulnerable to diarrhea and other water-borne diseases, so clean water led to declines in child mortality and, as a result, increases in life expectancy. Although history has endowed us with knowledge of the benefits of water disinfection, poor water quality and diarrheal diseases remain prevalent in many low- and middle-income countries today. Worldwide, diarrhea is the second leading cause of child mortality and a leading cause of morbidity. While this must partly be because of insufficient coverage of chlorination programs, it appears to be partly because the same disinfection technologies that were historically successful today tend to have weaker effects. It is therefore important to identify factors that limit realization of the full potential of water disinfection. In this study, we analyze average and local impacts of an abrupt, nationwide municipal water disinfection program that Mexico initiated in 1991, raising the share of the population receiving chlorinated water from 55% to 85% within six months. On average, the program was associated with a 37 to 48% decline in diarrheal disease deaths among children (over 23,000 averted deaths per year) and was highly cost-effective (about $1,310 per life year saved). However, age (degradation) of water pipes and the patchy coverage of complementary sanitation infrastructure play important roles in attenuating these benefits. Since, historically, infrastructural investments were made *pari passu* with chlorination efforts, this may explain their larger impacts. In our study, we also find some evidence of countervailing behavioral responses, insofar as private household expenditures on disinfectants decline in response to public program efforts, but these are small.

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Abstract

Historically, improvements municipal drinking water quality contributed significantly to mortality decline in wealthy countries. However, water disinfection has not produced equivalent benefits in developing countries today. We investigate this puzzle by analyzing a large-scale municipal water disinfection program in Mexico in 1991 that dramatically increased access to chlorinated water. On average, we find that the program led to a 37 to 48% decline in diarrheal mortality among children and was highly cost-effective ($1,310 per life-year saved). However, age (degradation) of water pipes and insufficient complementary sanitation infrastructure attenuated these benefits. Countervailing behavioral responses, although present, appear to be less important.

Keywords: clean water; chlorination; child mortality; infectious disease; diarrhea; Mexico; cost-effectiveness; sanitation; behavioral responses

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1. Introduction

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, Koppaka 2011, Preston and van de Walle 1978, Ogasawara and Inoue 2015, Ketzenbaum and Rosenthal 2014, Knutsson 2016). These disinfection technologies were often 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 (Turneaure and Russell 1901), diarrheal diseases due to poor water quality remain prevalent in many low- and middle-income countries today. Worldwide, diarrhea is the second leading cause of child mortality (Liu et al. 2012, Vos et al. 2015) as well as a leading cause of morbidity (roughly 1.7 billion episodes per year) (Fischer Walker, Rudan, et al. 2013). 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). It is therefore important to assess the effectiveness of water disinfection efforts in the developing world – and to identify factors that may limit realization of their full potential. With the exception of Galiani, Gertler, and Schargrodsky (2005), who analyze privatization of municipal water companies in Argentina, the recent literature on improving water quality in low- and middle-income countries has focused on interventions that are either small in scale or target rural populations with limited water system infrastructure (Arnold et al. 2013, Clasen and Haller 2008, Kremer et al. 2011, Zwane and Kremer 2007,

Duflo et al. 2015, Gruber et al. 2013, Arnold and Colford 2007). However, the urban population worldwide is projected to reach nearly 2.5 billion by 2050 (United Nations Population Division 2014), and the rapid growth of cities will require larger-scale interventions focused on municipal water supply (Brown, Keath, and Wong 2009, Fotso et al. 2007, McDonald et al. 2014, United Nations Development Programme (UNDP) 2006).

Improving urban water quality in developing countries on a large scale 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 (Shaheed et al. 2014, Lee and Schwab 2005). Low or intermittent water pressure in pipe networks can further reduce the efficacy of chlorination by increasing the degree of re-contamination (Kumpel and Nelson 2013, Ercumen et al. 2015, Jeandron et al. 2015, Long and Bliss 1937, 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 (Cutler and Miller 2006, United Nations Development Programme (UNDP) 2006, Gadgil 1998, Masten 2011, Crocker and Masten 2002). Second, improvements in water quality may not be effective without complementary investments in sanitation (Alsan and Goldin 2015, Duflo et al. 2015). Third, the provision of improved water may weaken private incentives for protective health behaviors. Such compensatory behavior holds the

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[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.

potential to crowd-out the health benefits of water disinfection (Bennett 2012, Keskin, Shastry, and Willis 2015).

Given these difficulties, it is not surprising that further large-scale investments in municipal water disinfection have received little emphasis in low- and middle-income countries, and that there are few studies of such interventions. A central contribution of this paper is that it studies a notable exception: *Programa Agua Limpia* (henceforth, PAL), a nationwide initiative to chlorinate existing municipal water systems across Mexico. PAL was launched in 1991 in response to a cholera epidemic that swept rapidly through Central and South America (Gutierrez et al. 1996, Sepulveda, Valdespino, and Garcia-Garcia 2006, Sepulveda et al. 2007). Within eighteen months, the number of localities disinfecting their water increased dramatically, from 250 to 15,000, and the share of Mexico's population receiving disinfected water rose from 55% to 90% (CONAGUA 1994). Importantly, this coverage was achieved without expanding existing piped water infrastructure or sewage networks.

To estimate the impact of PAL on child mortality across Mexico, we use detailed mortality statistics by age and cause at both the municipality and state level. While there was little variation in the timing of PAL implementation across regions, we are able to distinguish trend breaks in waterborne disease from potentially correlated omitted trends by using event study and difference-in-difference approaches in which the "control diseases" are those not directly influenced by water quality. To strengthen the identification of PAL program effects, we also estimate triple difference specifications, comparing changes in waterborne relative to control diseases across small-/middle-sized towns versus large cities (population>500,000). This specification exploits the fact that PAL targeted small- and medium-sized towns given that water in larger cities was already chlorinated. Importantly, in doing so, it accounts for other minor cholera prevention and treatment activities that were common across smaller towns and

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large cities – including legal restrictions on wastewater irrigation, hygiene campaigns, provision of Oral Rehydration Therapy (ORT), and general disease surveillance (CONAGUA 1994, Sepulveda, Valdespino, and Garcia-Garcia 2006).[2]

These different approaches yield similar results. We find that PAL was associated with a 36-48% reduction in diarrheal disease mortality rates among children under age 5, amounting to nearly 6,000 averted deaths per year. These reductions were entirely concentrated in the targeted small- and medium-sized cities. PAL was also highly cost-effective, with a cost per life year saved of about $1,310 (2015 USD). These substantive results are robust to the use of different transforms of mortality rates, mortality counts rather than rates, different event study windows, the inclusion of state-year fixed effects, and state- rather than municipal-level data.

Motivated by the contrasting experience of wealthy countries in history and developing countries today, we then study municipal-level variation in program effects (across more than 2,000 municipalities) 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 municipalities with greater baseline piped water coverage. We posit that this result is consistent with engineering concerns about recontamination through degraded, aged infrastructure (Mazari-Hiriart et al. 2005, Lee and Schwab 2005, Tulchinsky et al. 2000), and we provide suggestive direct 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

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[2] In practice water system chlorination was the overwhelming focus of the program (CONAGUA 1994).

literature (Alsan and Goldin 2015, Duflo et al. 2015), although it stands in contrast with null results found in the epidemiology literature (Fewtrell et al. 2005).[3]

Finally, we also use detailed household survey data to study behavioral responses to water system disinfection, revealing evidence of compensatory private behavioral responses among households. In particular, households in states that benefitted more from PAL reduced their spending on bottled water, soaps, and detergents – private health investments that may substitute for clean piped water (Bennett 2012, Keskin, Shastry, and Willis 2015). Notably, these reductions in household spending occurred only among households in small and medium-sized cities – precisely the areas targeted by PAL. Although we are unable to separately identify the direct health benefits of clean water from the indirect health decrements due to these behavioral responses, the degree of behavioral crowd-out is far from complete.

Overall, our findings contribute new evidence on the potential for large-scale public-sector water disinfection programs to improve health in developing countries. This represents a departure from the recent literature, which has either focused on smaller-scale interventions (Ahuja, Kremer, and Zwane 2010) or the role of privatization (Galiani, Gertler, and Schargrodsky 2005). In particular, our results suggest that such interventions indeed can be successful if they are supported by robust piped water infrastructure. We also find that their success may require concomitant investments in sanitation, consistent with emerging work in the economics literature (Alsan and Goldin 2015, Duflo et al. 2015). Finally, we also contribute to the literature on private behavioral responses to large public investments – and the extent to which behavioral responses

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[3] Recent experimental work from Tangiers shows that piped water adoption had no impact on health (or productive activities and income) because there was no improvement in water quality (Devoto et al. 2012). This illustrates the complementarity of water quality interventions with piped water provision.

undermine the benefits of these investments, both in the water sector and more generally (Bennett 2012, Das et al. 2013, Keskin, Shastry, and Willis 2015).

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, and Section 6 examines behavioral responses to program impacts. Section 7 presents estimates of the cost-effectiveness of PAL, and Section 8 then concludes.

2. 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 region of Mexico (*Figure 1*), with rates in small and medium sized towns nearly twice as high as those in cities.[4]

In 1991, a cholera pandemic emerged in Chile and Peru and quickly spread through South and Central America (Medina 1991, Sepulveda, Valdespino, and Garcia-Garcia 2006, Ries et al. 1992). 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, Valdespino, and Garcia-Garcia 2006). PAL

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[4] 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.

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).[5]

Despite these multiple components, 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 towns (the term "localities" in the Mexican census comprises all settlements regardless of size). In contrast, other (minor) program components were not specifically targeted to these areas and implemented more broadly (Sepulveda, Valdespino, and Garcia-Garcia 2006). Within 6 months of PAL's launch, the fraction of Mexico's population with chlorinated piped drinking water rose from 55% (almost exactly the percentage of the population living in large metropolitan areas) to 85%. By 1994, this figure exceeded 90% (*Figure 2*).[6]

Achieving these gains required a vast expansion of municipal water disinfection 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

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[5] Prior to the cholera pandemic, Mexico's primary approach to controlling diarrheal diseases emphasized treatment with Oral Rehydration Therapy (ORT), which we discuss below, and clinical case management (Frenk et al. 2003)

[6] Because implementation began in medium-size cities and then expanded to smaller towns, the increase in population coverage between 1991 and 1992 was much steeper than the corresponding increase in number of localities treated. Beginning in 1993, population coverage rose more slowly, but the number of localities treated continued to rise at the previous rate.

addition, to improve water quality in areas without piped water coverage, chloride tablets were disseminated to households (Sepulveda, Valdespino, and Garcia-Garcia 2006) 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.[7]

Prior work on PAL hints at large program effects. 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. (2004) show that morbidity declined by over 63% during the period 1990-1995. However, these studies essentially only describe national trends over time.

3. Data

3.1 Data

Our analysis uses data from three sources: the Mexican Vital Statistics, the 10% sample of the 1990 Mexican Population Census, and the *Encuesta Nacional*

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[7] 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. Similarly, efforts to modify hygienic behavior included use of radio and television to disseminate messages about hand-washing and other sanitary practices (Sepulveda, Valdespino, and Garcia-Garcia 2006). 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, Kremer, and Zwane 2010, Dupas 2011).

de Ingresos y Gastos de los Hogares (National Survey of Income and Expenditure, ENIGH).

      Mexican Vital Statistics. We obtained mortality data for neonates (age 0-1 months), postneonates (1-12 months), and children (ages 1-4) from the Mexican Ministry of Health. The vital statistics contain individual-level records of every certified death in the country.[8] 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).

      We aggregate individual-level deaths for children under the age of 5 into municipality-cause-year cells for all municipalities and years between 1985 and 1995. 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).[9] To convert diarrheal death counts into child (under-5), post-neonatal (1-12 months), and neonatal (0-1 month) mortality rates, we used data on the number of live births in each municipality and year.[10] We also repeat our estimation using death counts instead of death rates, finding similar results (see Section 4.1 below).

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[8] 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).

[9] 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.

[10] We obtained municipality-year data on live births from the *Instituto Nacional de Estadistica y Geographia* (INEGI, http://www3.inegi.org.mx/sistemas/microdatos/encuestas.aspx?c=33388&s=est). 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).

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To reduce the influence of extreme values (some municipalities are very small and have no recorded infant deaths and/or births in some years, for example), we trim municipalities in the top and bottom 5% of the average pre-intervention (1985-1990) diarrheal disease mortality rate distribution. This yields a sample of 2,223 of 2,438 municipalities, accounting for over 94% of Mexico's total under-5 population. We confirm that these sample restrictions do not alter our main findings.

A note about the quality of Mexico's vital statistics is also 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, Tome et al. 1997, Lozano-Ascencio 2008).[11] 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, Gonzalez and Quast 2011, Foster, Gutierrez, and Kumar 2009, Cutler et al. 2002). We cannot rule out the possibility that some degree of underreporting remains, but for this to bias our results, underreporting and/or misclassification must have changed sharply in 1991, and differentially for diarrheal diseases relative to control diseases.[12]

1990 Mexican Population Census. To analyze heterogeneous effects of PAL across municipalities, we use data on pre-program municipality

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[11] 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 mortality rates are under-counts and, in general, it is difficult to sign any resulting bias in the rates.

[12] The direction of any bias 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.

characteristics from a 10% sample of Mexico's 1990 population census (Minnesota Population Center 2015). Specifically, using appropriate sample weights, we aggregate household-level measures to create municipal-level data on the fraction of households with piped water and the fraction of households with septic system connections in their dwelling (a measure of sanitation coverage). We also 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 of piped water system age.

Encuestas Nacional de Ingresos y Gastos de los Hogares (National Survey of Income and Expenditure, ENIGH). To study behavioral responses to PAL, we use the ENIGH, a nationally representative income and expenditure survey. These data include detailed information about expenditures on bottled water, soaps, and detergents, which we combine into a single category. We use the 1989, 1992, and 1994 waves, which are the only waves falling within our study period.

Table 1 shows descriptive statistics for all variables used in our analyses.

3.2. Graphical Evidence

Figure 3 shows trends in child (under age 5) mortality rates from diarrheal diseases and "control" diseases between 1985 and 1995 using unadjusted data. We use two sets of control diseases to assess the sensitivity of our results to the choice of control illnesses. These are infectious respiratory diseases and non-respiratory childhood diseases other than diarrhea, respectively. The controls satisfy the condition that they are quantitatively important infectious causes of child mortality and are not *directly* influenced by water quality. We discuss the control diseases in more detail in Section 4.1.

There is a downward trend in diarrheal disease mortality rates among children under-5 prior to 1991, which is widely attributed to Mexico's Oral Rehydration Therapy (ORT) program that began in November 1984 (Mota-

Hernandez and Velasquez-Jones 1985, Gutierrez et al. 1996, Sepulveda, Valdespino, and Garcia-Garcia 2006). However, coincident with the implementation of PAL, in 1991, the decline becomes distinctly more rapid and pronounced. In contrast, there is no such trend break in mortality rates from the control diseases.[13],[14] Importantly, trends in diarrheal and respiratory disease mortality rates prior to PAL were quite similar. *Appendix Figure 1* shows trends over a longer period of time (1979-1997), and again the trend break in diarrhea-related mortality rates relative to control disease mortality rates remains clear.

4. Water Disinfection and Diarrheal Disease Mortality Rates
4.1. Empirical Strategy

      To formally estimate the relationship between PAL and mortality rates among children under-5 (altogether, as well as separately for neonates (0-1 month), post-neonates (1-12 months), and children ages 1-4), we test for differential trend breaks in diarrheal disease mortality rates in 1991 (when PAL was implemented) relative to two sets of control diseases which should not be directly affected by PAL.[15] One includes acute upper and lower respiratory

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[13] We formally test for a structural break between 1985-1995 in a longer series of data (1979-1997, *Appendix Figure 1*) using the Quandt Likelihood Ratio test (Quandt 1960). For diarrheal diseases, the test identifies a break in 1991, coincident with the start of PAL ($F = 11.16$, p = 0.004). We do not find evidence of a statistically significant break over the same period for both sets of control diseases (the identified break year is 1987, but this is not statistically significant - $F = 0.58$, p = 0.46 for non-respiratory controls; $F = 2.02$, p = 0.17 for respiratory controls).

[14] Notably, in the years just prior to PAL, the narrowing gap between diarrheal and non-diarrheal disease deaths appears to be driven by a slight *rise* in control disease death rates. We speculate that this slight increase may be related to Mexico's 1989-1990 measles epidemic. Mexico experienced a significant rise in measles attack rates as part of a broader regional pandemic in 1989-1990 (Katz et al. 2004). While we do not include measles mortality in either of our control disease sets, it is still possible that measles may account for the 1989-1990 uptick in these diseases given its association with respiratory diseases, congenital conditions, and skin infections (Orenstein, Perry, and Halsey 2004).

[15] A similar strategy was employed by Jayachandran, Lleras-Muney, and Smith (2010)) in their work on antibiotic therapy in the United States. Closer to our study, Galiani, Gertler, and Schargrodsky (2005)) use non-diarrheal causes of mortality as a falsification test in their study of water service privatization in Argentina.

infections, including viral bronchitis and pneumonia. We choose these diseases because they were the second-leading cause of child mortality in Mexico prior to PAL[16] 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 sensitive to *water quantity* (which affords opportunities for preventive handwashing and thus breaking the droplet contamination cycle), while diarrheal diseases are sensitive to *water quality*, which is the focus of PAL (Ahuja, Kremer, and Zwane 2010) – poor drinking water quality is not a direct risk factor for respiratory infections (Fischer Walker, Rudan, et al. 2013).

Nonetheless, if diarrheal diseases weaken immune systems, raising the likelihood of contracting respiratory infections (Fischer Walker, Perin, et al. 2013, Ashraf et al. 2013, Sedgwick and MacNutt 1910), then using respiratory diseases as control diseases will lead us to underestimate changes in mortality rates associated with PAL. For this reason, we also present results using an alternate set of control diseases that are less likely to co-vary with diarrhea (i.e., are bio-medically more 'distant').[17] These are perinatal causes (low birth weight, birth trauma, congenital infections), congenital anomalies, and a range of bacterial and viral illness (whooping cough, strep throat, scarlet fever, erysipelas, meningitis, tetanus, and septicemia) for which poor water quality is not a risk factor (Embrey et al. 2004, Pruss-Ustun et al. 2008).[18]

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[16] ICD-9 codes 460-466 and 480-487, respectively.

[17] Co-dependence between diarrheal diseases and these more bio-medically distant controls is also 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, Philipson, and Sala-i-Martin 1999). However, again, this would imply that our estimates are conservative.

[18] 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), and

We focus our analysis on municipality-year diarrheal mortality rates in a window around program implementation, 1985-1995 (*Appendix Table 2* shows that our substantive results are robust to expanding our study period). We begin by estimating event study specifications of the general form:

$$(1) \ M_{djt} = \alpha_d + \sum_{t=1985}^{1995} \alpha_t \left(\mathbf{1}(Diarrhea_d) \times \mathbf{1}(Year = t)\right) +$$
$$\sum_{t=1985}^{1995} \mu_t \left(\mathbf{1}(Year = t)\right) + \lambda_j + e_{djt}$$

Here, M_{djt} represents the inverse hyperbolic sine transformation of the mortality rate for disease class d in municipality (or state) j and year t.[19] $\mathbf{1}(Diarrhea)$ is a dummy variable denoting whether or not the cause of death is diarrheal disease (sensitive to water quality, and hence the program intervention) versus a control disease group (not sensitive to water quality); *Year* is a dummy variable denoting observations in year t; and λ_j represent municipality fixed effects. Estimates of α_t yield the average differential percentage change in the diarrheal disease mortality rate vs. the control disease mortality rate in year t relative to the baseline year (in this specification, 1985). We estimate this equation for overall under-5 mortality rates as well as for disaggregated age intervals (neonates 0-1 months, infants 1-12 months, and children 1-4 years old).[20]

"other bacterial diseases" (whooping cough, strep throat/scarlet fever, erysipelas, meningitis, tetanus, and septicemia; ICD9 codes 033-038).

[19] Formally, the inverse hyperbolic sine transform of some variable, y, is $\ln(y + \sqrt{(y^2 + 1)})$. Estimates obtained using this transformation can be interpreted in the same manner as those obtained using a natural logarithm transformation of the dependent variable, with the advantage of being defined at zero (Burbridge, Magee, and Robb 1988).

[20] This model effectively restricts pre-trends and breaks to be identical across each of the component control diseases which we sum together in forming our broader control disease groups. Thus, we additionally estimated versions of this model (as well as the more parametric formulations below), allowing for each of the component control diseases to have their own pre-trends, level, and trend breaks. This did not have any substantive impact on the findings.

To relax the assumption that the error terms are independent and identically distributed (i.i.d.) within municipalities over time, we consistently cluster the standard-errors at the municipality level (Bertrand, Duflo, and Mullainathan 2004). We also use municipality weights based on the average number of annual pre-intervention live births in all regressions.

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

$$(2) \ M_{djt} = \beta_0 + \beta_1 \ (\mathbf{1}(Diarrhea_d) \times \mathbf{1}(Post_t) \times Year_t)) + \beta_2 \ (\mathbf{1}(Diarrhea_d \times \mathbf{1}(Post_t)) + \beta_3 \ (\mathbf{1}(Diarrhea_d) \times Year_t) + \beta_4 \ \mathbf{1}(Diarrhea_d) + \sum_{t=1985}^{1995} \mu_t \ (\mathbf{1}(Year = t)) + \lambda_j + e_{djt}$$

Here, $\mathbf{1}(Post)$ is a dummy variable for post-PAL years (1991 and later), and all other variables are defined as before. Together, the parameters β_1 and β_2, measuring trend and level breaks in 1991, isolate the change in diarrheal disease mortality rates relative to the control disease mortality rates associated with PAL. β_3 captures pre-existing trends in diarrheal disease mortality relative to control diseases; the year fixed effects flexibly absorb mortality trend breaks and intercept shifts common across all diseases associated with PAL implementation, as well as common background (or pre-intervention) mortality trends; and β_4 reflects time invariant differences between diarrheal and control disease mortality rates.

Other less intensive prevention and treatment measures were also introduced around 1991. These components, which were substantially smaller in scope than chlorination, included legal restrictions on wastewater irrigation, hygiene campaigns, some provision of Oral Rehydration Therapy (ORT), and

general disease surveillance (CONAGUA 1994, Sepulveda, Valdespino, and Garcia-Garcia 2006). We would therefore expect estimates in equations (1) and (2) to largely reflect benefits of chlorinated water. However, to isolate the role of chlorination further, we also estimate the following triple-difference specification:

(3) $M_{djt} = \gamma_0 + \gamma_1 \, (\mathbf{1}(Diarrhea_d) \times \mathbf{1}(Small_d) \times \mathbf{1}(Post_t)) +$
$\quad \gamma_2 \, (\mathbf{1}(Diarrhea_d) \times \mathbf{1}(Small_d) \times \mathbf{1}(Post_t) \times Year_t) +$
$\quad \sum_3^4 \gamma_i \, (2 \; Triple \; Interactions) + \sum_5^9 \gamma_i \, (5 \; Double \; Interactions) +$
$\quad \sum_{10}^{11} \gamma_i \, (2 \; Linear \; Terms) + \sum_{t=1985}^{1995} \mu_t \, (\mathbf{1}(Year = t)) + \lambda_j + e_{djt}$

where $\mathbf{1}(Small_d)$ is a dummy variable denoting municipalities encompassing or contained within 'cities' with a population less than 500,000 (as enumerated in the 1990 population census),[21] and all other variables are defined as before.[22] The key parameters of interest are γ_1 and γ_2, which capture level and trend breaks in diarrheal disease mortality rates relative to control disease mortality rates in small- and medium-size municipalities relative to large cities.[23] Importantly, because chlorination explicitly targeted small- and medium-sized cities, these estimates are more granular in distinguishing the effects of chlorination from any potential effects of other more minor PAL components (not differentially targeted

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[21] 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. Any municipality that had a non-zero fraction living in such towns was coded as a large area, and the small area dummy is simply the inverse. Of note, 52% of the enumerated 1990 census population lived in larger areas, which is similar to the percentage of individuals with access to chlorinated water pre-PAL (55%, see *Figure 2*).

[22] The additional triple interactions terms (with subscripts omitted) are $\mathbf{1}(Diarrhea) \times \mathbf{1}(Small) \times Year$ and $\mathbf{1}(Small) \times \mathbf{1}(Post) \times Year$. The additional double interactions are $\mathbf{1}(Diarrhea) \times \mathbf{1}(Small;$ $\mathbf{1}(Diarrhea) \times Year; \mathbf{1}(Diarrhea) \times \mathbf{1}(Post); \mathbf{1}(Small) \times Year;$ and $\mathbf{1}(Small) \times \mathbf{1}(Post)$. The single terms are $\mathbf{1}(Diarrhea)$ and $\mathbf{1}(Small)$. $\mathbf{1}(Year),$ and $\mathbf{1}(Post) \times Year$ are subsumed by year fixed effects.

[23] Note that these models control for the full set of disease-municipality size specific pre-existing trends.

by city size).

We also conduct several robustness tests. First, we estimate our models using the natural logarithm of mortality rates rather than the inverse hyperbolic sine transformation. Second, we introduce state-year fixed effects into our specification to flexibly control for any time-variant state-level interventions or processes that may bias our results. Third, we re-estimate our specifications using state-level (rather than municipal-level) data to assess the sensitivity of our results to accounting for *all* registered deaths and births as well as any spillovers across municipal areas.[24] Fourth, we estimate our models using death counts rather than mortality rates. Fifth, we repeat our estimation using a longer time window (1979-1997).

4.2. Results

Figure 4 plots coefficient estimates obtained by estimating Equation (1) for under-5, neonatal (0-1 months), post-neonatal (1-12 months), and child mortality rates (ages 1-4 years) by event year. Consistent with the unadjusted graphs (*Figure 3*), *Figure 4* shows a moderate decline in diarrheal disease mortality rates relative to both sets of control diseases beginning in 1985 (with the implementation of Mexico's ORT program) – and then a discrete acceleration with the implementation of PAL in 1991, continuing throughout the post-implementation years. These results highlight the importance of conditioning on disease-specific mortality rate trends and motivate the parameterization of Equation (2).

Table 2 then reports estimates from Equation (2) by age group. In each

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[24] As discussed in Section 3.1, we restricted our sample to 95% of municipalities in culling the data of implausible values. Aggregating Vital Statistics data to the state level allows us to incorporate any deaths and births that our sample selection choices left out. Because there were no zero cells at the state-year level, we employed the natural log transformation for the dependent variables. For this and all other state level models, we compute wild cluster-T bootstrap standard errors (Cameron, Gelbach, and Miller 2008) to account for the relatively small number of clusters $(n = 32)$.

age interval, we find statistically significant level and trend breaks in diarrheal disease mortality relative to both sets of control diseases, and these breaks coincide with the introduction of PAL. Specifically, across all models, the estimate for "$\mathbf{1}(Diarrhea) \times Year$" (third row) shows that diarrhea mortality was declining before PAL, and the coefficient on "$\mathbf{1}(Diarrhea) \times \mathbf{1}(Post) \times Year$" (second row) shows a near doubling of this rate after 1991.

Overall, *Table 2* shows a sharp divergence in mortality rate trends between diarrheal diseases and control diseases that began in 1991 following the rapid implementation of PAL. For under-5 mortality rates, specifications using the non-respiratory controls (Panel A, first column) imply a 37% reduction in diarrheal disease mortality rates (more than 23,500 deaths averted) by 1995. The corresponding reduction using respiratory diseases as controls (Panel B, first column) is 48%.[25] Estimates for neonatal, post-neonatal, and 1-4 mortality rates imply similar declines (columns 2-4, *Table 2*).[26]

Table 3 reports estimates from Equation (3), exploiting an additional difference between the targeted small- and medium-size municipalities relative to larger cities. The first two rows show triple-difference estimates of level and trend breaks (γ_1 and γ_2, respectively) associated with PAL. These estimates imply a 38% and 37% reduction in diarrheal mortality rates among children under age 5 in specifications using non-respiratory and respiratory controls, respectively. Notably, these estimates are nearly identical to the double-difference estimates in

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[25] To calculate this, we multiplied the coefficient on the trend break, $\mathbf{1}(Diarrhea) \times \mathbf{1}(Post) \times Year$, by 4 (as 1995 is four years after the start of PAL) and added this to the coefficient denoting the level break ($\mathbf{1}(Diarrhea) \times \mathbf{1}(Post)$). Estimates using state level data (*Appendix Table 2*) yield similar results for models using non-respiratory controls. For models using respiratory controls, we estimate a 40% decline in diarrhea attributable to PAL by 1995.

[26] Using non-respiratory controls, the estimates for imply 0-1 month olds imply 22% decline, a 41% decline for 1-12 month olds, and a 25% decline for 1-4 year olds (Panel A, second-fourth columns, respectively). Using respiratory disease controls, the respective declines for these ages groups was 41%, 31%, and 17% (Panel B, second-fourth columns). We note that it is not surprising that we find large declines for neonates (0-1 month olds), for whom infectious diarrheal diseases are in fact a major cause of death (Liu et al. 2012).

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Table 2, confirming our *a priori* contention that other elements of PAL were not significant contributors to the decline in childhood diarrheal death rates.

4.3. Extensions and Robustness

Our results in *Table 2* are robust to using the natural log of mortality rates instead of the inverse hyperbolic sine (*Appendix Table 1*); the inclusion of state-year fixed effects (*Appendix Table 2*); using state-level rather than municipal-level observations (*Appendix Table 3*); using death counts instead of death rates (using both the inverse hyperbolic sine transformation in OLS regressions and raw counts in negative binomial model – see *Appendix Table 4*); a longer study window (*Appendix Table 5*).[27] Across all variants of our analyses for the time window 1985-1995, our point estimates using non-respiratory control diseases range from 35% to 41%, and our point estimates using respiratory control diseases range from 42% to 48%. Estimates using a wider event window (*Appendix Table 6*) suggest larger reductions (69% and 83% using non-respiratory and respiratory controls, respectively), though this may reflect the greater difficulty in tightly controlling for pre-trends in longer time series.

5. Under What Circumstances is Municipal Water Disinfection Effective? Estimating Heterogeneous Program Effects

The previous section demonstrates large average treatment effects of PAL. Given mixed experiences with water system chlorination/disinfection in many developing 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 first assess heterogeneity by coverage of municipal water pipe infrastructure, which delivers drinking water to households (and, importantly, *chlorinated* drinking water under PAL). In general,

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[27] Models using the longer-time series employ death counts, given restrictions on data availability on live births in prior years.

we would expect to find a positive relationship between pipe network coverage and PAL program effects. We then considered the role of water system age, positing that older water systems may have been more degraded (and hence less effective in providing clean water), attenuating program effects. We also study potential complementarities between PAL chlorination efforts and sanitation infrastructure coverage.

5.1. Empirical Strategy

To examine heterogeneous program effects, we first estimate municipality-specific program effects (DID_j). To do so, we subtract the difference in diarrheal diseases mortality rates before and after PAL from the same difference in control disease mortality rates, accounting for the pre-PAL trend in both disease groups.[28]

We then relate these to municipality-level piped water coverage, piped water system age, and sewage infrastructure coverage using data from the 1990 population census. Specifically, we estimate regressions of the form:

$$(4) \quad DID_j = \delta_0 + \delta_1(\textit{Piped Water}_j) + \delta_2(\textit{Sewage}_j) + \delta_3(\textit{System Age}_j) + \boldsymbol{\delta}(\boldsymbol{X}_j) + \\ \pi_s + e_j$$

where *DID* is the municipality specific program effect; *Piped Water Coverage* and *Sewage Coverage* represents the fraction of households in the 1990 census reporting having a piped water and sewage connection, respectively; and *System*

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[28] Specifically, we first compute the inverse hyperbolic sine transform of mortality rates and then compute the pre-trend between 1985 and 1991. We then subtract the predicted values from the post-intervention (transformed) mortality rates for each municipality-disease group. We then subtract the pre-intervention means from the de-trended post-intervention mean. Note that this is similar to estimating Equation (2) for each municipality j and then subtracting and using estimates of β_{1j} and β_{2j} to compute the percent change in diarrheal mortality rates in the pre- and post-PAL periods (after removing changes in control disease mortality and disease-specific pre-trends). Our results are similar in magnitude (and statistically indistinguishable) even when we account for differences in municipality-disease pre-PAL trends.

age is a proxy for the age of the water pipe infrastructure. X_j denotes a vector of municipality specific characteristics (percentage of adults completing secondary schooling, percent indigenous population, and the natural logarithms of average household earnings and population), and π_s denotes state fixed effects. We control for these variables to account for possible unmeasured confounders that may jointly influence municipality-specific treatment effects and water and sanitation infrastructure variables.[29] We estimate specifications using both the full sample and a restricted sample of small- and medium-sized cities (to focus more precisely on areas in which chlorination was introduced).

Equation (4) models the differential effect of chlorination under PAL according to pre-intervention measures of infrastructure. The coefficient estimates can be interpreted as interactions between PAL and each pre-intervention attribute. This advantage of this approach is that it allows us to examine several interactions simultaneously without loss of power.

5.2. Results

Piped Water Coverage. Figure 5 presents the bivariate non-parametric relationship between municipality-specific effects of PAL and piped water coverage; *Table 4* reports corresponding parametric estimates, both for the full sample and a restricted sample of municipalities with fewer than 500,000 people (which were targeted by PAL). Overall, we find no meaningful relationship between piped water coverage and mortality decline under PAL. While the estimates in *Table 4* show the expected sign, they are small and not statistically different from zero. Moreover, the estimated magnitudes actually *decline* when

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[29] One such 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, Estevez, and Magaloni 2016). However, Fried and Venkataramani (2016)) find no relationship between PAL treatment effects and the pre-PAL municipality governor's political party.

restricting the sample to those municipalities targeted by PAL. These results may help to explain the discordance between the historical experiences of wealthy countries and those of contemporary developing countries. We investigate this next.

Water System Age and Degradation. One possible explanation for the insignificant relationship between piped water coverage and PAL program effect size is that more extensive municipal water systems are older (Office of Economic Cooperation and Development (OECD) 2006, Vasquez et al. 2009, Oswald Spring 2011). Water system age is correlated with the degree of infrastructure degradation in many countries (Lee and Schwab 2005, Moe and Rheingans 2006, Larsen et al. 2016). Consequently, age-related pipe breaks, intermittent water pressure, and resulting recontamination from the surrounding soil may have undermined the full potential health benefits of the PAL (Ercumen, Gruber, and Colford 2014, Kumpel and Nelson 2013, Mazari-Hiriart et al. 2005). A substantial literature on Mexico suggests that older municipal water systems are more prone to infrastructure failure and fecal contamination (Mazari-Hiriart et al. 2005, Lee and Schwab 2005, Adler 2015, World Bank 1994).[30]

Because systematic data on water system age, direct measures of degradation, and water pressure (including pressure fluctuations) are not generally available for any country in the world,[31] including Mexico, this hypothesis is

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[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] 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 (18) 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.

difficult to test directly. To do so, 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 – which is reasonable in the Mexican context (Secretaria de Medio Ambiente y Recursos Naturales 2013) - higher ratios imply older water systems. *Figure 6* shows that this measure of pre-existing piped water coverage was positively associated with pipe breaks per kilometer in 2005 for a sample of 18 municipal systems.[32]

We then assess the relationship between our measure of water system age and piped water coverage in 1990 – and ultimately, municipal-level PAL program effects. *Figure 7* shows that areas with older water systems did in fact have greater piped water coverage in 1990. *Figure 8* then plots the non-parametric relationship between our measure of water system age and municipality-specific PAL program effects. In general, municipalities with older water systems experience smaller declines in mortality under PAL, suggesting a potentially important role for water system age (and degradation) in limiting the benefits of water system chlorination. Columns 3 and 4 of *Table 4* show that this relationship persists in multivariate parametric regression models.[33] The estimates from column 4 imply that municipalities with the oldest water infrastructure experienced a 45 percentage point decrease in PAL impacts, accounting for all of the mean treatment effect (*Table 2*). Thus, the role of system age and subsequent degradation may be substantial, to the point of blunting the positive impacts of chlorination altogether.

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[32] *Appendix Table 7* presents estimates from a regression of the logged pipe break measure on the access ratio measure. We find that a 1 s.d. increase in the access measure is associated with a 97% increase in pipe breaks.

[33] The sample sizes for these regressions are smaller due to the small sampling frame (1.5%) in the 1960 IPUMS census micro data.

Complementarity with Sanitation Infrastructure. *Table 4* shows a positive relationship between sanitation coverage at the municipal level and mortality decline under PAL. This suggests complementarity between water disinfection and sanitation efforts, a pattern consistent with recent work on Massachusetts during the nineteenth century (Alsan and Goldin 2015) and experimental findings from contemporary India (Duflo et al. 2015). Using the estimates in the second column of *Table 4*, an interquartile increase in municipality sewage coverage (from 9% to 49% of the population) was associated with a 31 percentage point greater decline in diarrheal disease mortality rates under of PAL. These results suggest that inadequate sanitation could also substantially limit the benefits water system disinfection in developing countries.

6. Behavioral Responses

In this section, we estimate behavioral responses of households to public investments in municipal water quality – another potentially important determinant of ultimate program impact (Bennett 2012, Dow, Philipson, and Sala-i-Martin 1999, Keskin, Shastry, and Willis 2015).

6.1. Empirical Strategy

Because the ENIGH surveys are nationally representative repeated cross-sections that do not sample the same municipalities over time, we study behavioral responses to PAL at the state-level using the following approach:

$$(5) \quad Y_{ijt} = \alpha_0 + \alpha_1 \left(\mathbf{1}(Post_t) \times Base\_Diarrhea_d\right) + \alpha_2 \left(\mathbf{1}(Post_t) \times Base\_Control_d\right)$$
$$+ \, \boldsymbol{\alpha} \times X_{ijt} + \lambda_i + \theta_t + u_{ijt,}$$

where Y_{ijt} is the inverse hyperbolic sine transform of total spending on mineral/purified water and soaps/detergents in household i, state j, and time t; *Base_Diarrhea$_d$* is the average under-5 diarrheal disease mortality rate during the

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four years prior to PAL; $\mathbf{1}(Post_t)$ is a binary indicator $= 1$ if the survey year was after PAL was initiated (1991 or thereafter); X_{ijt} is a vector of individual and household characteristics (age of household head and the natural logarithm of household income); and λ_j and θ_t represent state and year fixed effects (respectively).[34] Negative estimates of α_1, the key parameter of interest, would reflect compensatory behavioral responses under PAL. Although we do not have direct measures of hygienic behaviors, our assumption, which is supported by the epidemiology literature (Kamm et al. 2014, Pickering et al. 2010), is that household spending on disinfectants and clean water from alternative sources is a reasonable marker for sanitary behaviors that reduce diarrheal disease risk.

6.2. Results

Table 5 reports estimates from Equation (4). In our full sample, we find small, negative, and imprecise estimates of compensatory behavioral responses (estimates of α_1 shown in column 1). However, among households living in small and medium cities, the areas targeted by PAL, we find a statistically significant negative relationship with household spending on soap, disinfectants, and bottled water (column 2). Specifically, moving from the worst performing quartile of

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[34] This strategy was introduced by Card (1992)) and has recently been used in a number of health-related studies (Bhalotra and Venkataramani 2015b, Bleakley 2007, 2010, Cutler et al. 2010). Identification in this "continuous differences-in-differences" model comes from exploiting the sudden introduction of PAL (captured by *Post*) combined with variation across states in potential benefits from the program. The premise is that program impacts are larger in regions in which the disease burden is larger prior to introduction of the program, or that the program generates convergence in disease rates. We formally test for cross-state convergence in diarrheal mortality graphically and by estimating the following:

$$M_{jt} = \alpha_0 + \alpha_1 \left(\mathbf{1}(Post_t) \times Base\_Diarrhea_j \right) + \lambda_j + \theta_t + u_{jt},$$

where $M_{ijt} =$ the state specific diarrhea mortality rate and the other variables are as defined for Model 5 in the main text. An estimate of $\alpha_1 > 0$ is consistent with convergence. This model assesses absolute convergence. Estimating the same model using the logarithm of the mortality rate provides evidence of *relative* convergence. We find evidence of absolute convergence in child diarrhea mortality across states (*Appendix Figure 2* and *Appendix Table 6,* columns 2 and 4), but not of relative convergence (*Appendix Table* 6, columns 1 and 3).

states in the pre-program distribution of diarrhea mortality rates to the top quartile is associated with a 6.7% decrease in spending on these goods.[35] We find no statistically significant behavioral responses among households in large cities which had disinfected water prior to PAL (column 3).[36]

These results provide evidence of significant compensatory behavioral responses to PAL (or "crowd-out"). However, although we are unable to distinguish the independent health decrement associated with them from the direct health benefits of PAL, the fact that we nonetheless find large reductions in diarrhea mortality rates under PAL implies that the degree of crowd-out is far from complete.

7. Cost Effectiveness

PAL cost approximately U.S. $1 billion over the period 1991-1995 or $1.86 billion in 2015 dollars.[37] We have shown that PAL was associated with large reductions in diarrheal disease mortality, but a natural question is whether this was worth the expense relative to alternative uses of these public funds. To calculate the cost per child death averted and the cost per life year saved, we use the estimates from *Table 2* for total under-5 mortality rates (using non-respiratory controls). For each year between 1991 and 1995, we calculate the differential reduction in diarrheal disease mortality rates using the estimates for the level shift (*Diarrhea×Post)* and trend break (*Diarrhea×Post×Year)*. We multiply these estimated rate changes by the number of live births in each year to recover the

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[35] This effect appears to be driven primarily by soap and detergent purchases.

[36] Further, we find that impacts within these small and medium sized towns and cities were only seen for households with *both* sanitation and piped water access and not for households with only piped access or neither amenity (few households have sanitation access without piped water access). While these estimates (available upon request) are imprecisely estimated, the pattern of coefficients is consistent with the complementarity between disinfection and sanitation noted in Section 5.

[37] Personal communication with Dr. Jaime Sepulveda, who was Mexico's Vice-Minister of Health in 1991-1994.

implied number of averted deaths. We then sum the implied number of averted deaths across years 1991-1995. Overall, the implied cost per death averted under PAL was $78,600 (*Table 6*). 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 $1,310.[38] Estimates using the upper 95% CI of our estimates imply a cost per life year saved of $1,792 (or $1,050 using the lower 95% CI).

8. Discussion and Conclusion

One of the United Nation's Sustainable Development Goals (SDGs) for 2030 is to "ensure availability and sustainable management of water and sanitation for all" (United Nations 2015). Achieving this goal will require the provision of clean water to nearly 1 billion people who currently lack it worldwide. While many water quality interventions in developing countries focus on smaller scale approaches, the ambitious SDGs call for large-scale programs. Such interventions will be particularly important in cities, where rapid urbanization will further challenge existing – and in many cases already dilapidated water and sanitation infrastructure (United Nations Development Programme (UNDP) 2006).

Mexico's *Programa Agua Limpia* provides an unusual contemporary opportunity to study the potential of municipal water system disinfection on a national scale. On average, we find that the program successfully reduced childhood diarrheal disease mortality rates by over 30%. Moreover, we find that

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[38] Our estimate may be even more conservative for two additional reasons. First, we do not account for the stream of benefits accruing to PAL efforts *after* 1995. We do not consider this period because there is no published cost data around maintenance and upkeep of PAL-era disinfection infrastructure. However, one can imagine that the present discount value of life years saved exceeds reasonable estimates of the stream of program costs. Second, we do not account for additional program benefits beyond life years saved. For example, other work demonstrates robust benefits of long-run exposure to infant exposure to PAL on cognitive outcomes and schooling for girls (Bhalotra and Venkataramani 2015a).

PAL was highly cost-effective, with a cost per life year saved of $1,300. However, while some areas experienced large reductions, others enjoyed little mortality decline. We also find evidence of compensatory reductions in household hygiene investments, but the degree of crowd-out is far from complete.

To identify additional investments that may be necessary for municipal water disinfection to realize its full potential, we also examine the conditions under which PAL was effective. Studying municipal-level variation across the entire country, we find evidence suggesting that water system age, and associated degradation, may play an important role (because recontamination is more likely in degraded systems with intermittent water pressure, for example). We also find evidence of meaningful complementarities between water disinfection and pre-existing sewage infrastructure. Maximizing the returns to large-scale water quality interventions may require concurrent investments in repairing degraded pipes and in sewage infrastructure. Overall, we emphasize the complexity of contemporary urban water and sanitation challenges in developing countries – and the need for more research on how to make water systems more productive.

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Figure 1: Pre-Intervention Under-5 Diarrheal Mortality Rates Across Mexican Municipalities (1985-1990 average)
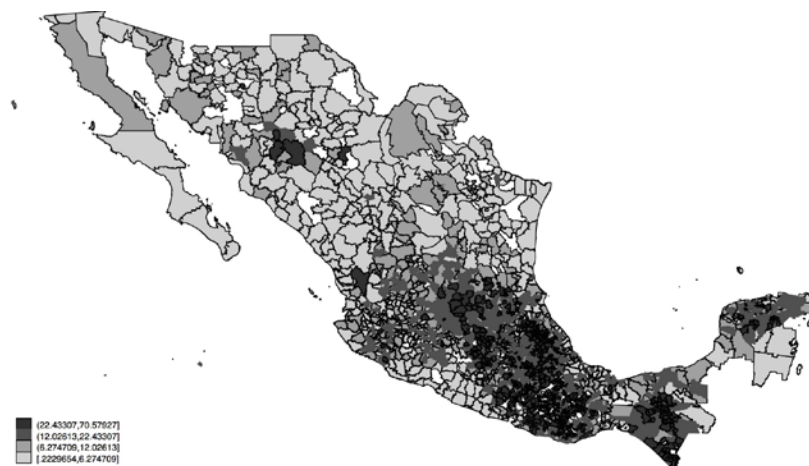
Notes: Map plots average diarrheal mortality rates per 1,000 live births over the period 1985-1990 for children under the age of 5 by municipality. Darker colors reflect higher average pre-intervention diarrheal mortality rates. Data to construct map were obtained from the Mexico Ministry of Health, Vital Statistics.

Figure 2: Chlorination Coverage under Programa Agua Limpia
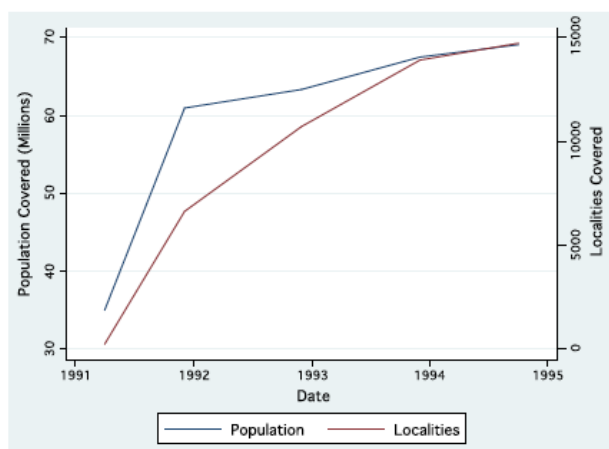
Notes: Figure plots the share of the population (blue) and number of localities (red) covered by chlorinated water between April 1991 and December 1995. The sharp uptick in the population share covered between April 1991 and December 1991 was coincident with the introduction of Programa Agua Limpia (PAL). Data were obtained from CONAGUA (1994).

Figure 3: National Trends in Log Under-5 Mortality Rates by Cause
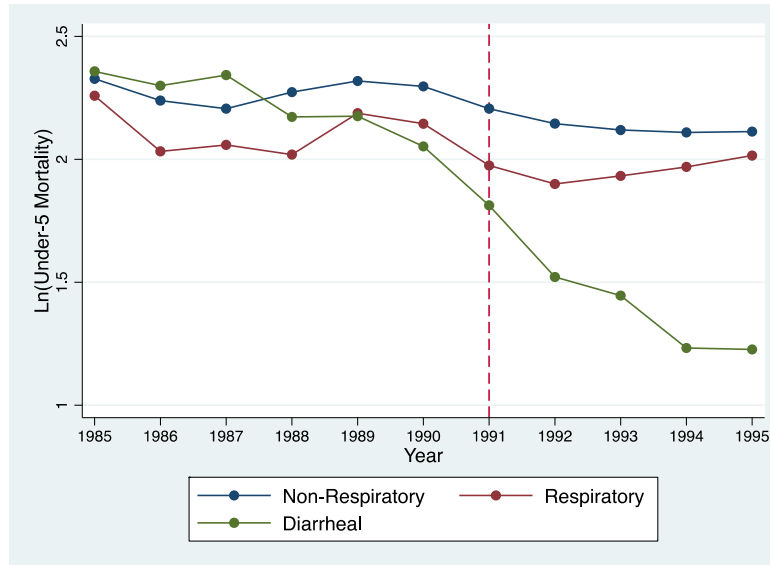

Notes: Figure plots the trends in mortality rates for diarrheal disease and the two sets of control diseases at the national level during the study period 1985-1995. Visually, we note a trend break in the diarrheal mortality series in 1991, coincident with the start of PAL. This break was confirmed econometrically using the Quandt Likelihood Ratio test ($F = 11.16$, p = 0.004) applied to a longer time series (1979-1997, see *Appendix Figure 1*; specifically, we assessed for the presence of breaks between 1985-1995). No statistically significant break was found for the control diseases. Data source: Ministry of Health, Vital Statistics.
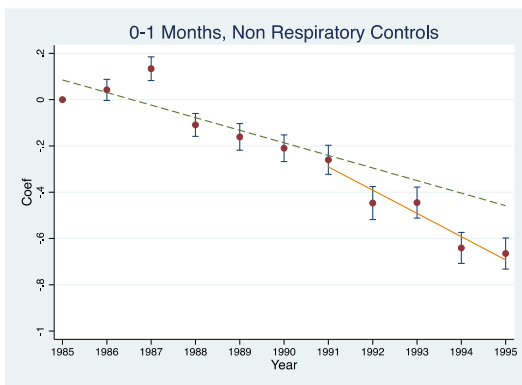
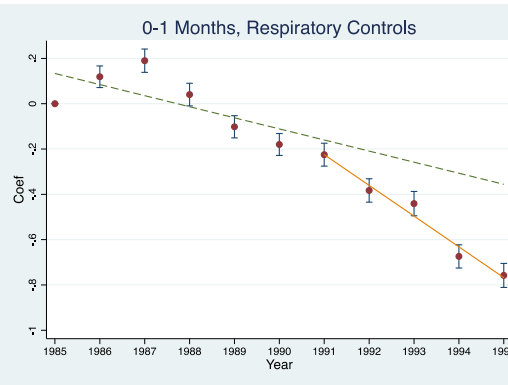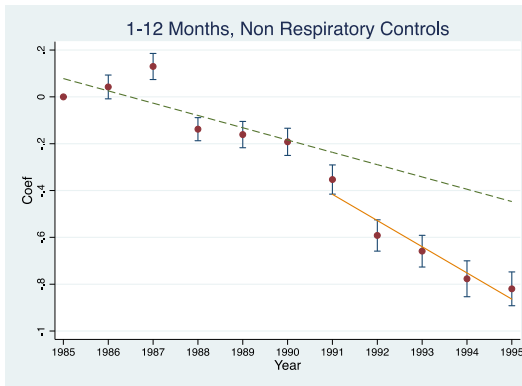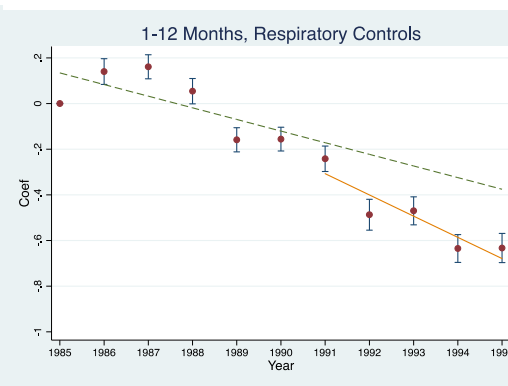**Figure 4: Event Study Coefficient Plots**

A
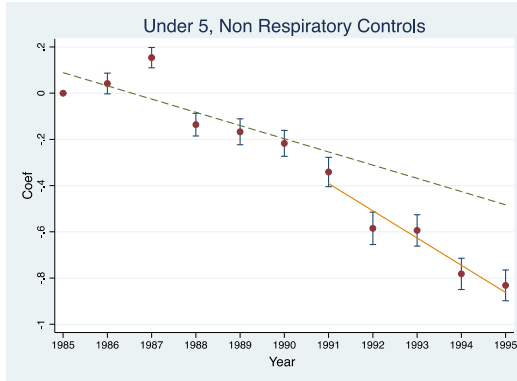
B
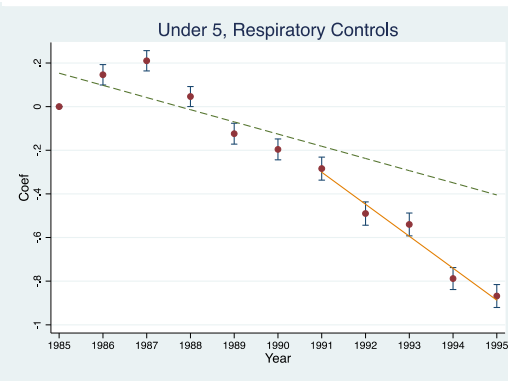
C

D
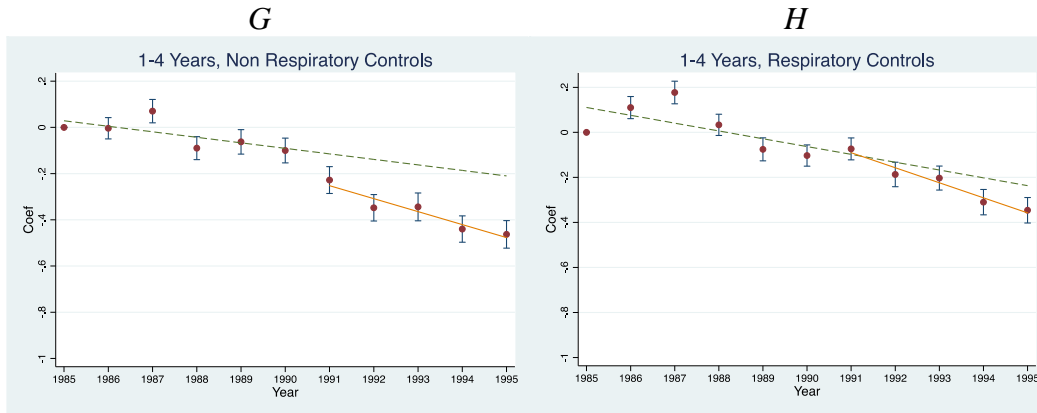

E


F

Notes: Panels A-H plot estimates of the Diarrhea × Year coefficients (α_t) from Equation (1) using either non-respiratory (left panel) or respiratory (right panel) control diseases. Panels A and B plot estimates for all children under age 5 and the remaining panels examine specific age subgroups (neonatal and post-neonatal refer to 0-1 month olds and 1-12 month olds, respectively). Coefficients are denoted by the dots and the vertical line and whiskers denote the 95% confidence interval of the estimates. The dashed green line is a predicted trend line calculated using pre-1991 coefficients, which we extend across the time series. The solid orange line is the trend line for the 1991-1995 coefficients. The gap between the two can be interpreted as the PAL treatment effect.

Figure 5: PAL Program Effects and Piped Water Coverage
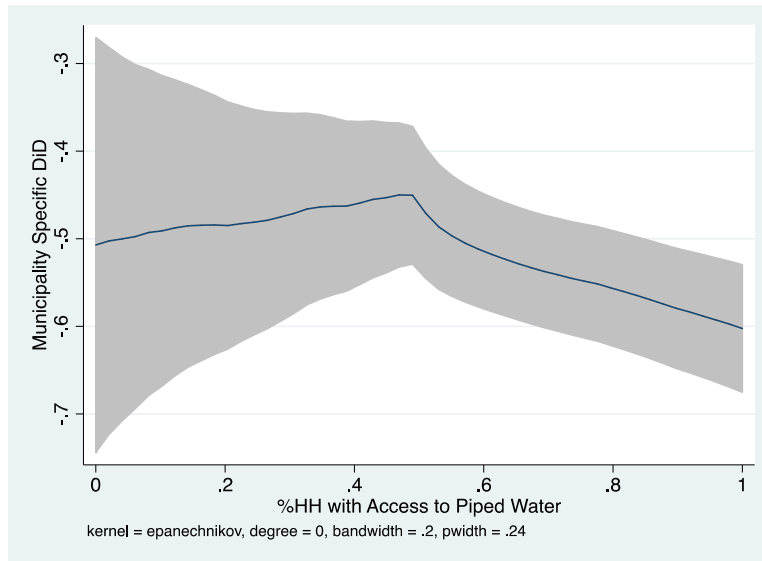
Notes: LOWESS smoothed plots (with 95% CI) of municipality specific differences-in-differences estimates (using respiratory disease controls) against the 1990 share of the population covered by piped water (see *Table 1* for details). More negative values on the Y-axis reflect larger program effects. The relationship plotted here is effectively flat.

40

Figure 6: Relationship Between Water System Age and Pipe Breaks
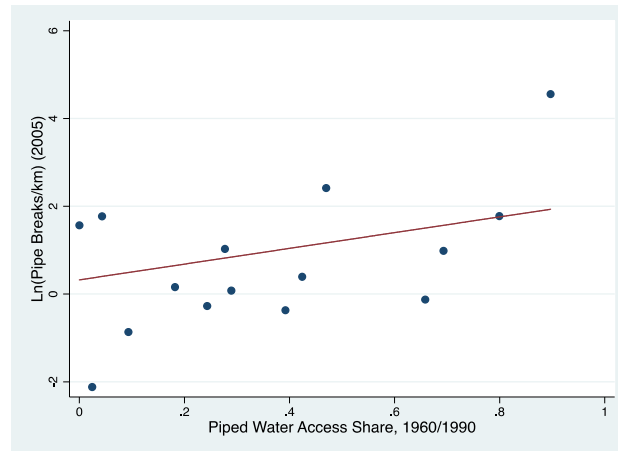
Notes: Scatter plot of logged pipe breaks per kilometer in 18 municipal water systems in 2005 against the ratio of households with access to piped water in (1960/1990), which we use as a proxy for system age. Data on pipe breaks were obtained from the International Benchmarking Network for Water and Sanitation Utilities (IB-NET, http://www.ib-net.org/). The regression line reveals a positive relationship between these variables. As seen in *Appendix Table 7*, a 1 s.d. increase in the piped water access ratio (0.21) is associated with a 97% increase in the number of pipe breaks per kilometer.

Figure 7: 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) against 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 6* directly makes the case for the latter point, linking the 1960/1990 access ratio to pipe breaks in 2005.
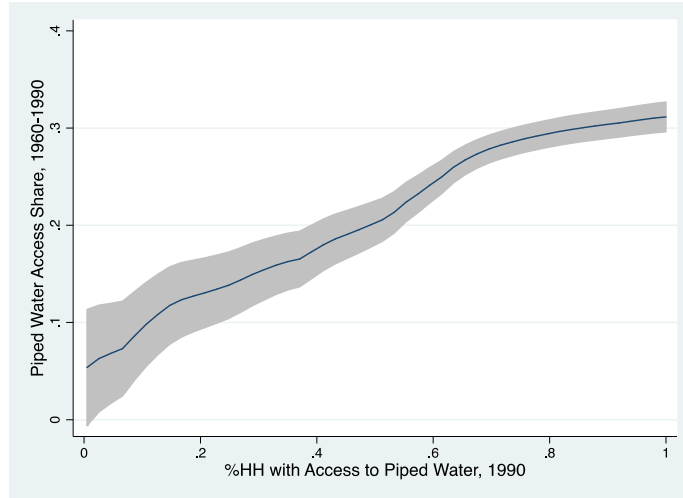
Figure 8: PAL Program Effects and Water System Age
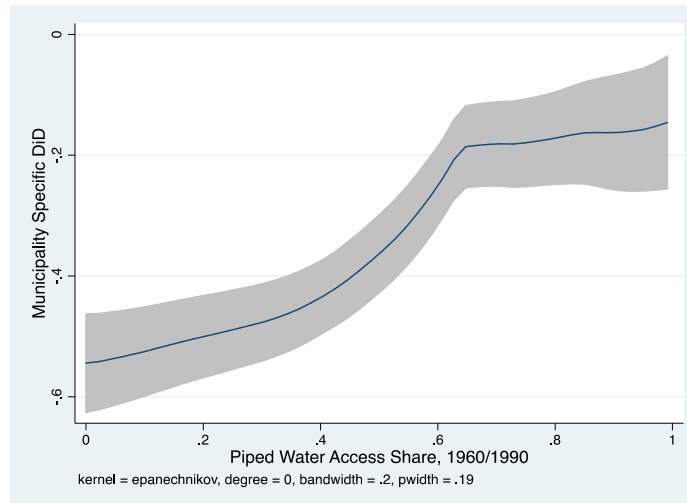
Notes: LOWESS smoothed plots (with 95% CI) of municipality specific differences-in-differences estimates and the ratio of piped water coverage between 1960 and 1990 (our proxy for water system age).

Table 1: Descriptive Statistics

	Pre-PAL			Post-PAL			Full Sample		
	Mean	S.D.	N	Mean	S.D.	N	Mean	S.D.	N
Vital Statistics Data (1985-1995, Municipality Level)									
Under-5 Diarrheal Mortality Rate	11.85	12.04	13,310	5.92	7.42	11,115	9.12	10.59	24,425
Control Diseases									
Under-5 Non-Respiratory Mortality Rate	10.5	10.49	13,310	9.22	7.96	11,115	9.90	9.33	24,425
Under-5 Respiratory Mortality Rate	11.01	10.19	13,310	9.24	7.41	11,115	10.19	9.60	24,425
Municipality Characteristics, 1990 Census									
% HH with Access to Piped Water	0.79	0.21	2,331						
% HH with Access to Sewage	0.60	0.29	2,331						
% Adults Completing Secondary Schooling	5.35	1.57	2,331						
Average Per Capita Earnings (Pesos, 1000)	42529.42	17434.71	2,331						
% Indigenous Population									
Population	348,272.9	1,627,442	2,331						
ENIGH Data (1989, 1992, 1994 waves; Individual Level)									
Total Expenditure on Sanitation Goods (Peso)	1428	1113.15	11,176	810.11	1360.57	22,558	1015.022	1316.469	33,734
Household Head Age	45.05	15.43	11,176	44.58	15.39	22,558	44.74	15.41	33,734
Household Head Gender (=1 if Male)	0.86	0.35	11,176	0.86	0.34	22,558	0.86	0.34	33,734
Household Head Less then Primary (=1)	0.49	0.50	11,176	0.51	0.50	22,558	0.50	0.50	33,734
Primary Complete (=1)	0.25	0.43	11,176	0.24	0.43	22,558	0.24	0.43	33,734
Secondary Complete	0.12	0.33	11,176	0.13	0.34	22,558	0.13	0.33	33,734
Diploma or Above	0.15	0.35	11,176	0.13	0.34	22,558	0.13	0.34	33,734
Household with Piped Water (=1)	0.60	0.50	11,176	0.53	0.50	22,558	0.55	0.50	33,734

Notes: Descriptive statistics divided by pre-PAL (data from prior to 1991) and post-PAL (data from 1991 or thereafter). Vital Statistics data reflect weighted under-5 deaths per 1000 live births, with the sample size denoting municipality-year observations. Diarrheal diseases denote infectious diarrhea (ICD9 codes 001-009). Non-respiratory diseases comprises of the sum of death rates from congenital anomalies (ICD9 codes 740-759), perinatal causes (low birth weight, birth trauma, congenital infections, neonatal jaundice, etc: ICD9 codes 764-779), and other bacterial diseases (whooping cough, strep throat/scarlet fever, erysipelas, meningitis, tetanus, and septicemia; ICD9 codes 033-038). Respiratory diseases consist of the sum of death rates from acute upper and lower respiratory infections (ICD9 codes 60-466, 480-487 respectively). Municipality characteristics 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. 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). Finally, the ENIGH data are taken from the 1989, 1992, and 1994 waves of the *Encuestas Nacional de Ingresos y Gastos de los Hogares*, a national household income and consumption survey. The sanitation goods variable reflects the total amount spent in the prior month on soaps, detergents, and bottled or purified water. We combine these into a single class to reduce the number of zeros, since we used logged expenditures in our empirical models.

Table 2: Differences-in-Differences Models: Diarrheal Relative to Control Disease Trend Breaks Pre-Post PAL

	Panel A - Non Respiratory Controls				Panel B - Respiratory Controls			
	Under-5 Years	0-1 Months	1-12 Months	1-4 Years	Under-5 Years	0-1 Months	1-12 Months	1-4 Years
1(Diarrhea)*1(Post)	-0.138	-0.048	-0.181	-0.136	-0.119	-0.0651	-0.136	0.00773
	(0.0210)	(0.0208)	(0.0232)	(0.0199)	(0.0204)	(0.0201)	(0.0208)	(0.0199)
1(Diarrhea)*1(Post)*Year	-0.0572	-0.0419	-0.0582	-0.0297	-0.0909	-0.0868	-0.0422	-0.0321
	(0.00867)	(0.00919)	(0.00906)	(0.00732)	(0.00824)	(0.00819)	(0.00832)	(0.00756)
1(Diarrhea)*Year	-0.0591	-0.0561	-0.0543	-0.0259	-0.0558	-0.0489	-0.0509	-0.0347
	(0.00543)	(0.00558)	(0.00561)	(0.00512)	(0.00467)	(0.00474)	(0.00488)	(0.00429)
1(Diarrhea)	-0.086	-0.643	1.057	0.601	-0.133	-0.361	0.147	0.144
	(0.0485)	(0.0523)	(0.0328)	(0.0294)	(0.0387)	(0.0401)	(0.0313)	(0.0246)
N	48,906	48,906	48,906	48,906	48,906	48,906	48,906	48,906
R-squared	0.52	0.62	0.51	0.52	0.53	0.50	0.50	0.48
% Decline by 1995 Due to PAL	36.7	21.6	41.3	25.4	48.2	41.2	30.5	17.1

Notes: Estimates of Equation (2) in the main text. Robust standard errors, correcting for clustering at the municipality level, are provided in parenthesis. All models include municipality and year fixed effects and are weighted by municipality baseline live births (i.e., the average for 1988-1990). Each column represents a separate regression, with the panel header denoting the control diseases used and the column header the specific age group over which the inverse hyperbolic sine transformation of mortality was calculated for the dependent variable. **1**(*Diarrhea*) =1 denotes diarrheal mortality rates, while **1**(*Diarrhea*) = 0 denotes the control disease (see Table 1 notes for details). *Post* = 1 if the year of observation is 1991 or thereafter. The level (**1**(*Diarrhea*) × **1**(*Post*)) and trend (**1**(*Diarrhea*)× **1**(*Post*) ×*Year*) break estimates are in bold. The coefficients on **1**(*Diarrhea*) ×*Year* pick up pre-existing trends in the diarrheal disease series relative to the control diseases. Pre-trends and level and trend breaks for the control diseases are picked up the year fixed effects

The sample size reflects the number of municipality-years. The final row, *% Decline by 1995 Due to PAL* is calculated by adding the coefficient on the level break to four times that on the trend break (since 1995 is 4 years after PAL). The year variable is rescaled such that 1991 = 0 and increments above and below are denoted by positive and negative integers.

Table 3: Triple Difference Models: Diarrheal Disease Mortality × Post-intervention × Small (Targeted) vs Large Municipality

	Non-Respiratory	Respiratory
1(Diarrhea)*1(Post)*1(Small)	0.0332	0.0693
	(0.0459)	(0.0460)
1(Diarrhea)*1(Post)*1(Small)*Year	-0.104	-0.109
	(0.0181)	(0.0194)
N	48,906	48,906
R-squared	0.553	0.566
% Decline by 1995 Due to PAL	38.3	36.8

Notes: Estimates of Equation (3) from the main text. Robust standard errors, correcting for clustering at the municipality level, are in parenthesis. All models include municipality and year fixed effects and are weighted by municipality baseline live births (i.e., the average of 1988-1990). Each column is a separate regression, and reflects estimates of Equation (3) in the main text. Specifically, we include a third difference - municipality size - in these models. Size is captured by a binary indicator denoting municipalities with populations < 500,000 = 1. Those larger already generally chlorinated water systems in place. While only 5% of municipalities in the sample are contained in large urban areas > 1,000,000 people (Small = 0), they accounted for 52% of the population in 1990.

The final row, % Decline by 1995 Due to PAL is calculated by adding the coefficient on the level break to four times that on the trend break (since 1995 is 4 years after PAL). The year variable is rescaled such that 1991 = 0 and increments above and below are denoted by positive and negative integers.

Table 4: Determinants of Municipality-Specific Program Effects

	Full Sample	Municipalities in Small-Medium Sized Cities	Full Sample	Municipalities in Small-Medium Sized Cities
% HH with Access to Piped Water	-0.374	-0.307	-0.232	-0.232
	(0.311)	(0.330)	(0.358)	(0.391)
% HH with Access to Sewage System	-0.785	-0.840	-1.01	-1.21
	(0.398)	(0.450)	(0.438)	(0.509)
Piped Water Coverage Ratio (1960/1990)			0.284	0.603
			(0.162)	(0.258)
N	1,921	1,837	1,408	1,334
R-squared	0.16	0.16	0.18	0.19

Notes: Estimates of Equation (4) in the main text. Robust standard errors are in parenthesis. All models are at the municipality level and are weighted by municipality population. The dependent variable is the municipality specific PAL treatment effect. We restrict the sample to the 5th-95th percentile of this variable in order to reduce the influence of outliers. The first and third columns include all observations meeting the stated criteria. In the second and fourth columns, we remove all municipalities within, or that contain, cities with >500,000 (see main text), as these areas already had high penetrance of chlorinated water prior to PAL. All models include municipality level controls from the 1990 census, including the % of population completing secondary education, average logged household income, % indigenous population, logged total municipality population, and state fixed effects.

Table 5: Behavioral Responses: Expenditures on Soaps, Detergents, Bottled Water

	Full Sample	Small/Medium Sized Cities	Large Cities
1(Post)*Base Diarrhea	-0.00847	-0.0167	-0.00253
	(0.0130)	(0.00739)	(0.0210)
	[0.60]	[0.13]	[0.85]
N	34,801	21,385	13,416
R-squared	0.874	0.887	0.839

Notes: Estimates of Equation (4) from the main text. Robust standard errors, correcting for clustering at the state level, are in parenthesis. Each column represents a separate regression, with the column header denoting the specific sub-sample of interest. Because there are only 32 states, we additionally compute cluster wild bootstrap-t corrected p-values for the coefficient of interest (Cameron et al., 2008), which are provided in the square brackets. The dependent variable is the inverse hyperbolic sine transformation of expenditures on soaps, detergents, and bottled/mineral water. All models include controls Post×Baseline Respiratory Disease, age, sex, and education of the household head, household size, and logged household total income, as well as state and survey rear FE. The main coefficient of interest is on **1**(Post) ×Base Diarrhea, where Baseline Diarrhea is the average under-5 diarrheal mortality rate over 1988-1990. To interpret the results, focusing on the second column an interquartile decrease in diarrheal mortality (about 4 deaths per 1,000 live births) was associated with a 6.7% decrease in soap, detergent, and bottled water expenditures.

Table 6: Cost Effectiveness

| Estimate	Under-5 Deaths Averted	Program Cost (2015 USD)	Cost/Death Averted (2015 USD)	Cost/Life Year Save (2015 USD)
Point	23664	1.86 billion	78600	1310
Lower 95% CI	17298	1.86 billion	107527	1792
Upper 95% CI	29535	1.86 billion	62976	1050

Notes: Assuming 2.5 million live births each year (the average nationwide between 1985-1990), we use estimates from Col 2 of Table 2 to calculate the number of under-5 deaths averted each year between 1991-1995. We do this using both the point estimates from the model, as well as the lower and upper 95% CI. The column "Under-5 Deaths Averted" displays these estimates. The total program cost over 1991-1994 comes from administrative sources. The cost per death averted is simply the total cost/# 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.

Urban Water Disinfection and Mortality Decline in Developing Countries

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Appendix Figure 1: National Trends in Under-5 Deaths By Cause, 1979-1997
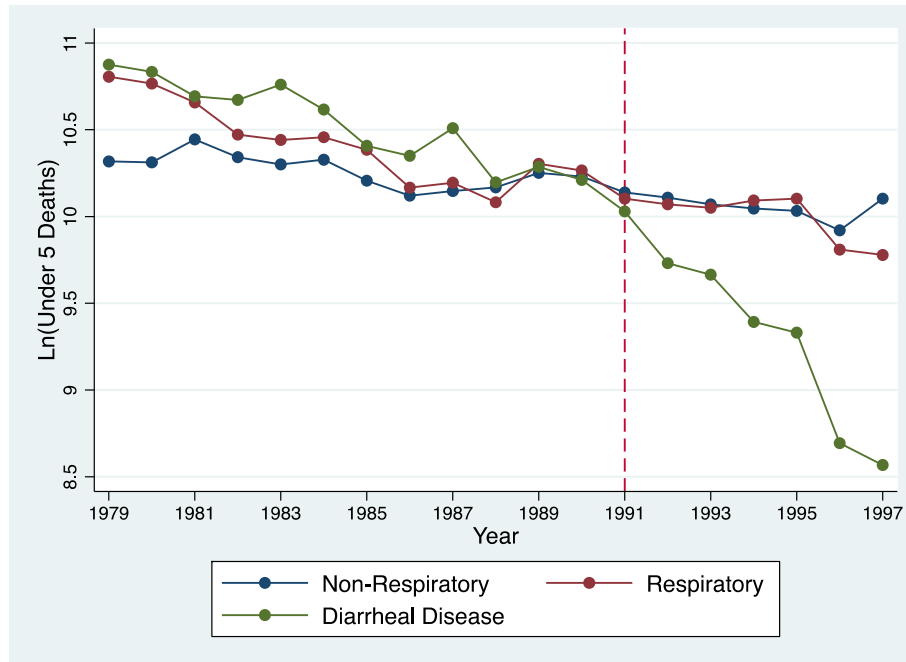
Source: Mexican Vital Statistics. Unlike *Figure 3*, we plot deaths instead of deaths rate here because municipality specific live births were not available for the entire period.

Appendix Figure 2: Convergence in Diarrheal Mortality Rates
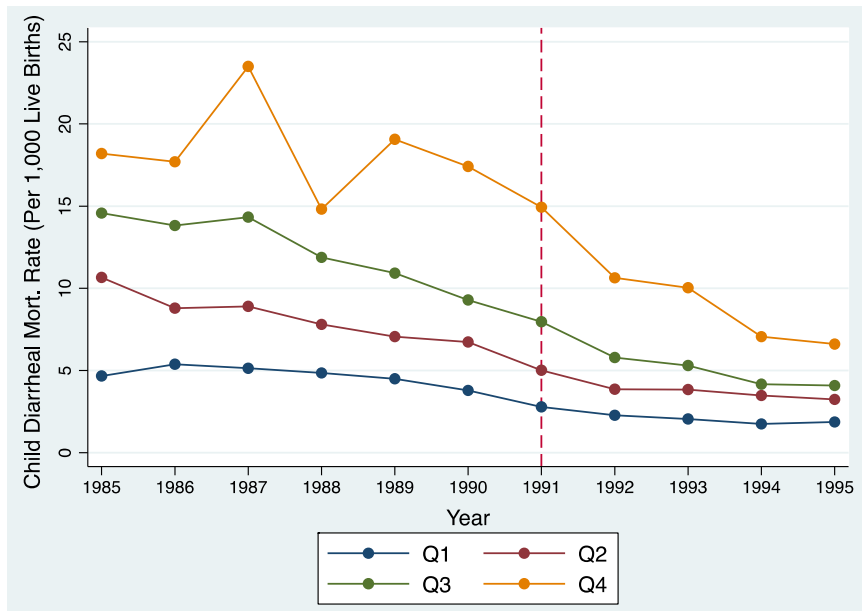
Source: Mexican Vital Statistics. Figure plots trends in under-5 diarrheal mortality rates by quintile of (state) pre-intervention diarrheal mortality (calculated as the average over 1988-1990). The figure demonstrates gradual convergence prior to 1991, with more rapid convergence thereafter.

Appendix Table 1: DID Trend Break Models: Logged Mortality Rates

	Non-Resp	Respiratory
1(Diarrhea)*1(Post)	-0.135	-0.108
	(0.024)	(0.023)
1(Diarrhea)*1(Post)*Year	-0.055	-0.078
	(0.009)	(0.010)
N	35,203	36,234
R-squared	0.59	0.64
% Decline by 1995 Due to PAL	35.7	42.3

Notes: Estimates of Equation (2) from the main text. Models are identical to those presented in *Table 2* except here we use natural logs of the mortality rates instead of the inverse hyperbolic sine transform. This accounts for the smaller sample sizes in these models. The final row, % Decline by 1995 Due to PAL is calculated by adding the coefficient on the level break to four times that on the trend break (since 1995 is 4 years after PAL). The year variable is rescaled such that $1991 = 0$ and increments above and below are denoted by positive and negative integers. See *Table 2* notes for further details.

Appendix Table 2: DID Trend Break Models: Inclusion of State-Year FE

	Non-Resp	Respiratory
1(Diarrhea)*1(Post)	-0.138	-0.119
	(0.0206)	(0.0200)
1(Diarrhea)*1(Post)*Year	-0.0572	-0.0909
	(0.00850)	(0.00808)
N	48,906	48,906
R-squared	0.594	0.563
% Decline by 1995 Due to PAL	36.7	48.3

Notes: Estimates of Equation (2) from the main text. Models are identical to those presented in *Table 2* except here we include state-year level fixed effects. The final row, % Decline by 1995 Due to PAL is calculated by adding the coefficient on the level break to four times that on the trend break (since 1995 is 4 years after PAL). The year variable is rescaled such that $1991 = 0$ and increments above and below are denoted by positive and negative integers. See *Table 2* notes for further details.

Appendix Table 3: DID Trend Break Models: State-Level Data

	Non-Resp	Respiratory
1(Diarrhea)*__**1**(Post)	-0.151	-0.112
	(0.038)	(0.0366)
	[p= 0.000]	[p=0.02]
1(Diarrhea)*__**1**(Post)*Year	-0.065	-0.0774
	(0.026)	(0.020)
	[p = 0.02]	[p=0.004]
N	698	698
R-squared	0.75	0.77
`		
% Decline by 1995 Due to PAL	41	42

Notes: This table is identical to Table 2 in the main text except here we use state-level data. All models include state and year fixed effects. The dependent variable is logged under-5 mortality. Robust standard errors, correcting for clustering at the state level, are in parenthesis. Because there are only 32 states, we additionally compute cluster wild bootstrap-t corrected p-values for the coefficient of interest (Cameron et al., 2008), which are provided in the square brackets.

Appendix Table 4: DID Trend Break Models: Monthly Death Counts

	OLS		Negative Binomial	
	Non-Resp	**Respiratory**	**Non-Resp**	**Respiratory**
1(Diarrhea)*1(Post)	-0.085	0.032	-0.065	0.038
	(0.032)	(0.033)	(0.034)	(0.036)
	[p=0.001]	[p=0.33]		
1(Diarrhea)*1(Post)*Year	-0.066	-0.119	-0.058	-0.116
	(0.020)	(0.018)	(0.021)	(0.021)
	[p =0.004]	[p = 0.002]		
N	8,335	8,335	8,376	8.376
R-squared	0.86	0.85		
% Decline by 1995 Due to PAL	34.9	44.4	31.2	40.3

Notes: This table is identical to Table 2 in the main text except here we use monthly data for all states between 1985-1995 (i.e, observations are aggregated to the state-month level). We additionally include month fixed effects in these models. Observations are aggregated to the state-month level. The dependent variable in the OLS models is the logged number of deaths in each state-month. The dependent variables in the Negative Binomial models is the number of deaths. For the negative binomial models, we use the incidence rate ratios (IRR) to calculate the percent relative decline in diarrheal deaths. Robust standard errors, correcting for clustering at the municipality level, in parenthesis. Cluster wild bootstrap-t corrected p-values (Cameron et al., 2008), are provided in the square brackets.

Appendix Table 5: DID Trend Break Models: Municipalities, 1979-1997

	Non-Resp	*Respiratory*
1(Diarrhea)* **1**(Post)	-0.0872	-0.243
	(0.0227)	(0.0218)
1(Diarrhea)* **1**(Post)*Year	-0.151	-0.148
	(0.00568)	(0.00535)
N	80,766	80,766
R-squared	0.87	0.87
% Decline by 1995 Due to PAL	69.1	83.5

Notes: Estimates of Equation (2) from the main text. This table is identical to *Table 2* in the main text except here we use data from 1979-1997, instead of restricting the sample to 1985-1995. The dependent variable is the inverse hyperbolic sine transform of the number of deaths under age 5 (we do not examine mortality rates since municipality specific births data was not available during the early part of the period). Robust standard errors, correcting for clustering at the municipality level, are in parenthesis.

Appendix Table 6: Cross-State Convergence

	(1)	(2)	(3)	(4)
	Log	**Level**	**Log**	**Level**
1(Post)*Base Diarrhea	**0.00758**	**–0.328**	**0.0312**	**–0.0480**
	(0.0135)	**(0.116)**	**(0.0123)**	**(0.0302)**
Base Diarrhea*Year			0.00461	0.0115
			(0.00193)	(0.0172)
1(Post)*Base Diarrhea*Year			**–0.0137**	**–0.0409**
			(0.00258)	**(0.0165)**
1(Post)			-0.507	-0.270
			(0.0839)	(0.111)
Year			-0.0924	-0.184
			(0.0123)	(0.0506)
1(Post)*Year			-0.0202	0.142
			(0.0173)	(0.0469)
N	349	352	349	352
R-squared	0.96	0.777	0.941	0.830

Notes: Estimates provided are versions of the model:

$$M_{jt} = \alpha_0 + \alpha_1\ (Post_t \times Base\_Diarrhea_j) + \lambda_j + \theta_t + u_{jt}$$

where M_{jt} = the state (j) specific mortality rate, *Base_diarrhea* refers to the baseline under-5 diarrheal mortality rate (average for each state over 1988-1990), *Post = 1* for years 1991 and thereafter, and the remainder of the variables reflect state and year fixed effects. Estimating using logged mortality rates as the dependent variable evidence of *relative* convergence, where as estimates using levels recover *absolute* convergence. A negative coefficient on post*base would implies convergence. The 1985-1995 state-level sample is used for all regressions (N reflects the number of state-years).

In Cols 1 and 2, we find a precisely estimated 0 for logged mortality and a negative and significant estimate for the level, implying that there was absolute, not relative, convergence. Columns 3 and 4 decompose this convergence into a level and trend break. Again, we find strong evidence of absolute (level) convergence. While we obtain statistically significant estimates on the level and trend break for logged mortality in col 3, these effects are offsetting such that by 3 years after program initiation, the net effect on mortality rates is zero.

Robust standard errors, correcting for clustering at the state level, in parenthesis. Statistical significance is unchanged when we compute cluster wild bootstrap-t corrected p-values for the coefficient of interest (which we suppress here for purposes of clarity).

Appendix Table 7: Water System Age and Pipe Breaks	
	ln(Pipe Breaks/Km
Piped Water Access Ratio, 1960/1990	4.64
	(0.99)
% change from 1 s.d. increase in access ratio	*97%*
N	18
R-squared	0.69

Notes: Estimates from model regressing logged pipe breaks per kilometer in 2005 on the ratio of households with access to piped water in 1960/1990 (weighting by 1990 population). Data on pipe breaks were obtained from the International Benchmarking Network for Water and Sanitation Utilities (IB-NET, http://www.ib-net.org/). Percent change estimates reflects the % increase in pipe breaks owing to a 1 s.d. change in the piped water access ratio (0.21).