Local Human Capital Development and Firm Resilience: Evidence from the 1997 Indonesian Crisis

Parth Chawla*

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Abstract

Do returns to human capital rise during crises? This paper examines whether human capital accumulation generated by Indonesia's INPRES school construction program in the 1970s improved firm resilience during the 1997 Asian Financial Crisis. I find that each additional school per 1,000 children increased post-crisis real labor productivity and output by 2.8 and 3.5 percent, respectively. INPRES induced changes in local pre-crisis workforce composition that shifted workers toward more skill-intensive production work. I argue that these skills became particularly valuable during the crisis, when firms faced disruptions. Consistent with this mechanism, INPRES had larger effects in sectors with higher import intensity, where adjustment demands were greater. Finally, I show that the local abundance of skilled workers in high-INPRES districts kept wages lower, enabling plants to retain more of them when they were most valuable.

JEL classification: O12, J24, D22, G01

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*Department of Agricultural and Resource Economics, University of California, Davis, 1 Shields Ave, Davis, CA 95616. Email: chawla@ucdavis.edu.

1 Introduction

> "[T]he ability to deal successfully with economic disequilibria is enhanced by education ... this ability is one of the major benefits of education accruing to people privately in a modernizing economy." (Schultz, 1975)

There is a premium on human capital during economic crises. To deal with the disequilibrium of a crisis, firms must reallocate resources, adjust production decisions, and adapt to new economic conditions while minimizing losses. Their ability to navigate the crisis successfully depends on the speed and effectiveness of these decisions. Firms often respond by laying off workers to cut costs, securing additional financing to maintain cash flow, or implementing technological and organizational changes (Guimbert and Oostendorp, 2016; Kwapil, 2010; Westergård-Nielsen and Neamtu, 2012). Since workers are responsible for making and executing these decisions, the strength of a firm's crisis response is inevitably shaped by their skills.

The role of worker characteristics in firm performance during recessions has received little attention in the economics literature. The standard view, dating back to Schumpeter (1939), is that the strongest, most efficient firms survive. A natural extension of this idea is that highly skilled workers contribute to this survival by helping firms adapt. Indeed, following Schultz's (1975) argument, one could argue that skills matter the *most* during periods of crises, as that is when they are tested to their limits. The impact of a plant manager's decision to optimize production, the speed and precision with which a machine operator adapts to new technology, and the ability of a procurement specialist to secure critical raw materials amid supply chain disruptions all become even more important in a crisis. When firms are on the brink of shutdown, these factors can determine whether they adapt and survive or fail.

In this paper, I examine the impact of Indonesia's INPRES school construction program on human capital accumulation and its role in facilitating firm recovery after the 1997 Asian Financial Crisis. Between 1973 and 1979, the Indonesian government undertook an ambitious expansion of the education system, dou-

bling the stock of schools by constructing 61,800 new primary schools across the country. This large-scale intervention significantly improved educational attainment. Duflo (2001) found that the program increased schooling levels by 0.25 to 0.40 years, with each additional school per 1,000 children contributing 0.12 to 0.19 years of education.

In a later paper, Duflo (2004) examined the medium-run labor market effects of the INPRES program from 1986 to 1999. By 1986, thirteen years after the program began, the first cohort fully exposed to INPRES schools had turned 18 and begun entering the workforce, raising the overall share of workers with primary education. By the time the 1997 crisis struck, individuals from regencies (referred to as "districts" hereafter) with more INPRES schools had disproportionately benefited from the program and made up a growing share of the workforce, as older, less-educated workers exited. This provides a natural setting to assess whether improved human capital made firms more resilient to the crisis.

Despite exhibiting similar pre-crisis trends, plants in above-median INPRES-intensity districts were less adversely affected by the 1997 crisis than those in below-median-intensity districts. I attribute this pattern to what I call the human capital salience hypothesis, namely, that the returns to human capital increase during periods of crisis. Under this view, the human capital accumulated through INPRES became especially valuable when plants faced disruptions and had to rapidly implement protective or adaptive responses. This, in turn, improved firm resilience during the crisis.

To test this hypothesis, I use data from the Statistik Industri (SI) census of manufacturing plants with 20 or more employees from 1987 to 1999, combined with data from Duflo (2001) on INPRES intensity at the district level. Using a difference-in-differences approach, I find that INPRES significantly improved post-crisis plant performance. On average, each additional INPRES school per 1,000 children in a plant's district led to a 2.8 percent increase in labor productivity (real output per worker) and a 3.5 percent increase in real output after 1997, compared to what would have been expected otherwise among surviving plants. These effects are larger in specifications estimated on the full SI census with district fixed effects, compared

with within-plant estimates that include plant fixed effects.

Importantly, re-estimating the main specification with plant exit (defined as dropping out of the SI census) as the outcome yields no statistically significant effects. Consistent with Hallward-Driemeier and Rijkers (2013), this reflects a weak relationship between productivity and exit during the crisis period. Together with the within-plant results, this suggests that the estimated resilience effects are not driven by compositional change or selective attrition. In particular, I find no evidence of differential selection by INPRES intensity. That is, the productivity-exit relationship weakens after 1997, but not differentially in high-INPRES districts.

I next assess the robustness of the estimated resilience effect to a range of alternative explanations and sources of confounding. I find no evidence of differential pre-trends between high and low-INPRES exposure districts prior to the crisis. Because INPRES targeted areas with lower school enrollment that were typically poorer and more rural, I examine whether the results reflect pre-existing differences in local conditions that predate the program. The results remain robust to controls for pre-crisis measures of local capital intensity and financial dependence, as well as to controls for district-level pre-crisis in-migration rates of INPRES-exposed cohorts, addressing concerns about pre-crisis sorting of INPRES-educated workers across districts. I also find no differential impact by plant size or capital intensity, indicating that the results are not driven by larger or more capital-intensive plants having greater scope for adjustment. The effects are likewise not driven by foreign ownership, government ownership, or group affiliation.

The results are also robust to using a balanced panel of 4,895 plants observed in every year from 1990 to 1999 rather than the full SI sample, excluding pre-crisis outliers in labor productivity to ensure that the estimates are not driven by a small number of plants, applying plant-level price deflators rather than a national GDP deflator to account for sectoral, regional, and local price variation during the crisis, and including sector-by-year fixed effects.

Plants in high-INPRES districts outperform particularly in sectors with high import intensity, where pro-

duction is more exposed to supply-chain disruptions and requires more complex adjustment. This pattern is consistent with the human capital salience hypothesis. Skills do not manifest as faster growth in normal times but become more valuable when firms must reoptimize inputs and production processes during the crisis. Interestingly, INPRES effects are weaker among exporters, which is plausibly consistent with exporters benefiting from the large rupiah devaluation during the crisis.

Because these effects cannot be explained by simple differences in basic educational attainment, I further assess whether they reflect broader human capital gains by examining how INPRES affected occupational composition. Using population census data and the SAKERNAS labor force survey, I find that INPRES increased the probability of employment in skilled production-related occupations, defined primarily as ISCO-68 major groups 7 and 8 (craft and related trades; plant and machine operators and assemblers), along with a small number of similar occupations from other groups. This pre-crisis shift toward skill-intensive production work is consistent with broader human capital improvements that may have supported greater resilience during the downturn.

In the mechanisms section, I provide further evidence that human capital is the operative channel behind the observed effects. Beyond raising schooling, INPRES may have expanded the local stock of higher-quality human capital in other ways in districts where more primary schools were built. Newer schools may have offered better infrastructure or instruction, and early access to schooling may have raised aspirations, encouraging further vocational preparation. By the 1990s, a growing share of the workforce consisted of younger cohorts who had attended INPRES schools, increasing the relevance of these human capital gains during the crisis.

In addition to the pre-crisis shift toward skilled production occupations, I find that plants in high-INPRES districts were more likely to have workers engaged in training before the crisis. During the crisis, plants with greater exposure to INPRES were more likely to retain or expand their workforce, and the workers they retained were more likely to have completed primary education. Finally, although wages declined

across all worker types, the gap between nonproduction and production wages widened, driven primarily by a sharper decline in production worker wages.

To explain how these patterns contribute to crisis resilience, I outline a simple model in which the variable inputs to production are high and low-human capital labor. A negative demand shock reduces the marginal product of labor, but the decline is smaller for high-human capital workers, whose skills are more complementary to the capital and organizational adjustments that plants undertake in response to the shock (the human capital salience hypothesis). High-INPRES plants benefit from an abundant local supply of skilled production workers, which keeps their wages relatively low despite their higher marginal product. This allows plants to retain more skilled workers and sustain output during the downturn. Finally, I provide additional supporting evidence for my human capital salience hypothesis by showing that the productivity premium to primary education increased during the crisis.

This paper contributes to several strands of the economics literature. First, it contributes to the literature on the direct, indirect, and long-term impacts of the INPRES program, and large-scale investments in educational infrastructure more broadly. Following the seminal work by Duflo (2001, 2004), several papers have examined its effects, such as Martinez-Bravo (2017), who found that INPRES increased the education levels of local village heads, leading to greater provision of public goods. More recently, Akresh, Halim and Kleemans (2022) studied its long-term and intergenerational effects on outcomes such as employment sector, migration, marriage outcomes, and the education and living standards of women and their children. Hsiao (2024) also showed that labor mobility greatly amplified the wage gains from the program.

Second, it contributes to the literature on firm risk-coping strategies, resilience, and performance during economic crises, a topic that has received much attention in management research but remains understudied in development contexts. Most notably, Bartoloni, Arrighetti, and Landini (2021) examine firm survival probabilities conditional on firms adopting corporate strategies that emphasize skill accumulation. They find that Italian manufacturing firms that invested in human capital, innovation, or international expansion

had higher survival rates following the 2008 Great Recession.

In the Indonesian context, Hallward-Driemeier and Rijkers (2013) use the same dataset as this paper to show that the 1997 crisis led to the exit of relatively productive firms but also led to the entry of more productive ones, partially offsetting the crisis's negative impact on aggregate productivity. While their findings suggest that the crisis weakened the link between productivity and firm survival, my results highlight how pre-crisis skill accumulation helped firms better withstand the shock. Narjoko and Hill (2007) also use the same dataset and find that foreign ownership, export orientation, and industry labor intensity were positively associated with firm survival after the crisis.

Finally, this paper closely relates to the literature on household resilience to shocks (Barrett et al., 2021). Recent studies, including Do, Nguyen and Grote (2025), D'Errico, Romano and Pietrelli (2018), and Smith and Frankenberger (2018), show that human capital and education plays a significant role in enhancing households' adaptive capacity to negative shocks.

2 Context and Data

2.1 The 1997 Crisis in Indonesia

Indonesia experienced rapid industrial growth, averaging over 9 percent annually from the 1970s until the onset of the 1997 crisis (Narjoko and Hill, 2007), making it one of the fastest-growing economies in Southeast Asia. The country maintained low inflation, modest current account deficits, rapid export growth, and strong foreign exchange reserves. Given these conditions, many economists expected Indonesia to be relatively insulated from the crisis triggered by the devaluation of the Thai baht in July 1997 (Iriana and Sjöholm, 2002). However, the contagion spread quickly, and Indonesia became one of the hardest-hit countries, facing massive capital outflows, a sharp depreciation of the rupiah, surging inflation, and a severe economic contraction. GDP declined by 13 percent in 1998. Manufacturing was hit especially hard due to its reliance

on imported inputs and exposure to exchange rate volatility (Hallward-Driemeier and Rijkers, 2013).

2.2 Plant Data

I use plant-level data from the Statistik Industri (SI) census of manufacturing plants with 20 or more employees from 1987 to 1999. This is administered by Indonesia's statistical agency, Badan Pusat Statistik (BPS). The data include detailed information on plant output, employment, investment, capital stock, material inputs, wages, and ownership. Importantly for my analysis, the dataset also provides information on worker education levels from 1995 to 1997, which allows me to test underlying channels and mechanisms.

The dataset provides broad coverage across regions and sectors, with plants classified using Klasifikasi Lapangan Usaha Indonesia (KLUI-1990) codes, which are largely compatible with ISIC Revision 2. The data allow me to control for 5-digit sector and district fixed effects in regressions and also enable a within-plant specification with plant fixed effects, as plants are observed repeatedly over time. However, a limitation of the data is that it does not distinguish between firms and plants, meaning I cannot identify whether multiple plants are owned by the same entity.

Since district boundaries in Indonesia have changed frequently over this period, I harmonize them to their 1995 definitions to merge with the INPRES data. Additionally, the dataset includes product-level price and quantity information from 1994 onward, allowing me to test robustness to controlling for price changes by using plant-level deflators instead of a GDP deflator, following Cali et al. (2024, 2022).

2.3 INPRES, Census, and Demographic Data

To measure regional INPRES intensity, I use district-level data from Duflo (2001) on the rollout of the Sekolah Dasar INPRES primary school construction program. Launched in 1973, this large-scale initiative aimed to expand access to primary education by rapidly increasing the number of schools. Between 1973 and 1979, the government built 61,800 new schools, effectively doubling the number of schools per capita.

Figure A.1 in Appendix A shows regional variation in program intensity and the spatial distribution of plants in the SI sample. Districts with lower initial enrollment rates in 1972 received more schools.

For my shift-share strategy, I combine the district-level share of INPRES schools constructed with a time-varying national share of working-age adults (ages 18-60) born in 1968 or later (the cohorts fully exposed to the program). These national exposure shares are calculated using single-age population data from the UN World Population Prospects.

To examine pre-crisis occupational changes associated with INPRES, I combine the 1990 Population Census with the 1995 Intercensal Population Survey (SUPAS), both from BPS and accessed through IPUMS International. In addition, I use 1990 census microdata to construct district-level pre-crisis in-migration rates for the INPRES-exposed cohort, also obtained from the IPUMS International database.

3 Descriptive Statistics

Table 1 shows summary statistics for key variables. There was substantial cross-district variation in the scale of the INPRES program. On average, district-level INPRES intensity is 2.34 schools per 1,000 children. When expressed as a share of total INPRES schools, each of the 290 districts received about 0.35 percent of schools on average. This version is used as the "share" component in the shift-share (Bartik) instrument for the 2SLS regressions. The "shift" component, which provides time variation in treatment exposure, is the national share of the workforce born in or after 1968; this share rises from 4 percent in 1986 to 47 percent in 1999.

The next part of the table summarizes the plant-level census data from 1987 to 1999. The average manufacturing plant employs 165 workers and produces output worth approx. 524 million rupiah (in 1995 prices), though there is wide variation in scale. One plant, likely a data error, reports 116,052 workers; the next largest has 42,649. Labor productivity, measured as the log of output per worker (in thousands), averages 9.07 with a standard deviation of 1.33, reflecting substantial productivity differences across plants.

Worker education data, available for the 1995-1997 subsample, show that 91 percent completed at least primary school, 52 percent junior high, 28 percent senior high, and only 1 percent attained tertiary education.

Table 1: Summary Statistics from INPRES, Census and Plant Data

	Obs	Mean	Std. Dev.	Min	Max
District INPRES intensity (schools per 1000 kids)	290	2.33505	1.25235	0.59082	8.59827
District share of INPRES schools (share)	290	0.00345	0.00185	0.00087	0.01270
National share of the workforce born $\geq$ 1968 (shift)	14	0.27356	0.13758	0.04297	0.47183

	Obs	Mean	Std. Dev.	Min	Max
Total workers	161555	165.04	679.37	10.00	116052.00
Log total output	161555	13.17	2.04	3.40	23.14
Log output per worker	161555	9.07	1.33	-3.51	17.30
Log value added	161553	12.07	1.98	-80.78	22.44
Log total investment	80465	11.86	2.84	0.34	24.27
Log total machinery value	23735	11.12	2.53	0.69	21.19
Log total wages per worker	161414	7.01	0.88	-0.24	14.30
Share workers primary and above	51409	0.91	0.19	0.00	1.00
Share workers junior high and above	51409	0.52	0.34	0.00	1.00
Share workers senior high and above	51409	0.28	0.29	0.00	1.00
Share workers bachelors and above	51409	0.01	0.03	0.00	0.59
Shift $\times$ Share (plant-level)	161555	0.00084	0.00051	0.00007	0.00599

Plant data cover manufacturing plants from 1987 to 1999.
Monetary variables are expressed in thousands of 1995 rupiah; logarithms are natural logs.
"Shift × Share" is the shift-share variable used as the instrument in the 2SLS regressions.

Figure 1 provides an initial look at the time series of the main plant outcomes before and after the crisis. Plants in above-median INPRES districts consistently lag behind their below-median counterparts in both output and labor productivity, though the trends move in parallel. When the crisis hits, outcomes fall across the board, but the decline is noticeably sharper in below-median districts.

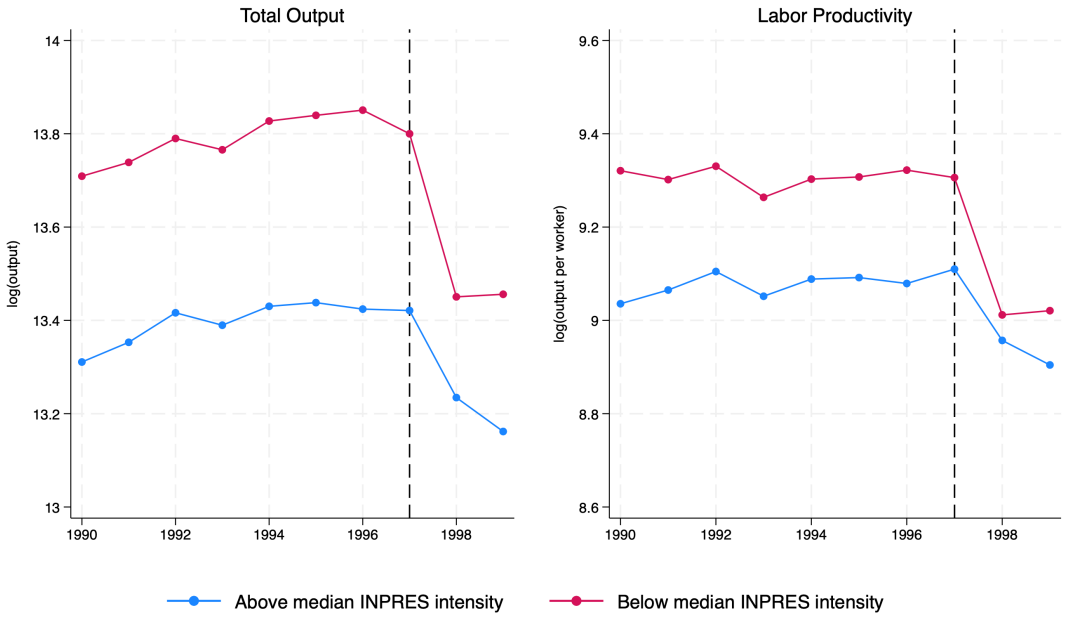
Overall, the descriptive evidence highlights three facts that motivate my empirical strategy in the next section:

(i) INPRES intensity varies widely across districts and was predetermined by the late 1970s.

(ii) The national share of the workforce born after 1968, who benefited from INPRES, rises substantially over time.

(iii) In the simple means-over-time comparison, above-median INPRES districts show a smaller decline in outcomes during the 1997 crisis than below-median districts, despite similar pre-crisis trends.

Figure 1: Total Output and Labor Productivity by INPRES Intensity, 1990-1999



Notes: The figure shows yearly means of plant outcomes in districts above and below the median INPRES intensity, weighted by the number of plants per district-year. Labor productivity is real output per worker. The dashed line marks the 1997 Asian Financial Crisis. Monetary variables are expressed in thousands of 1995 rupiah; logarithms are natural logs.

4 Empirical Strategy

To test whether INPRES created resilience for plants during the 1997 crisis, I first describe a difference-in-differences framework that compares the change in plant outcomes before and after 1997 in districts with high INPRES intensity against the corresponding change in districts with low INPRES intensity.

$$y_{isdt} = \alpha + \beta(\text{Post 1997} \times \text{INPRES})_{dt} + \nu_s + \lambda_t + \phi_g + \varepsilon_{isdt}. \quad (1)$$

Here, y_{isdt} denotes either labor productivity (real output per worker, in logs) or real output (in logs) of plant i in sector s, located in district d at time t. The variable INPRES captures the intensity of the INPRES school construction program, defined as the number of schools built per 1,000 children in district d. ν_s, λ_t, and ϕ_g are sector, year, and unit fixed effects, respectively, where ϕ_g can be either district-specific or plant-specific fixed effects, depending on the specification.

The coefficient β represents the effect of an additional INPRES school on a plant's outcome during 1997-1999. The main source of variation comes from district-level differences in INPRES intensity, which was not randomly assigned. The school construction occurred in the 1970s, and individuals born in or after 1968 were fully exposed, as they entered primary school after the new schools were built. These cohorts began entering the labor force around 1986, at age 18. As successive post-1968 cohorts entered the labor market, the share of the workforce exposed to INPRES schools increased over time.

Each plant around 1997 may have employed a mix of younger workers, who were more likely to have attended INPRES schools, and older cohorts who were not exposed. Since worker-level exposure cannot be directly observed, I construct a shift-share (Bartik) instrument by multiplying the national share of the workforce (ages 18-60) born in or after 1968 with each district's fixed INPRES school share. Specifically, I instrument the main variable of interest from equation (1), $(\text{Post 1997} \times \text{INPRES})_{dt}$, with the following:

$$Z_{dt} = \big(\text{Post 1997}\big)_t \times \big(\text{District share of INPRES schools}\big)_d \times \big(\text{National share of workers born } \geq 1968\big)_t. \tag{2}$$

The baseline DiD term, $(\text{Post 1997} \times \text{INPRES})_{dt}$, is zero in every district before 1997 and, from 1997 onward, equals that district's INPRES intensity. Because the intensity itself does not change over time, all post-crisis observations carry the same treatment "dose". The Bartik instrument introduces additional time variation by multiplying each district's share of INPRES schools with the national share of fully treated cohorts (born $\geq$ 1968) in the working-age population (ages 18-60) for each post 1997 year, so the effective

treatment dose facing plants in every district rises steadily from 1997 to 1999.

Table 1 shows that the mean district share of INPRES schools is 0.00345, and the national share of workers born in or after 1968 rises from 0.418 in 1997 to 0.472 in 1999. Multiplying the fixed district share by this shifting cohort weight implies that Bartik exposure for each district's plants increases by about 13 percent over the crisis window, even though the underlying district-level INPRES intensity remains unchanged. I will later show that the impact of INPRES after 1997 becomes stronger in districts with higher exposure when interacted with these national cohort shifts (Table A.1 in the Appendix).

Although the cohort shifts introduce useful time variation in the treatment dose, they are few in number and common across all districts. As a result, identification still primarily relies on the cross-district variation in INPRES shares, consistent with the "exogenous-share" Bartik design described by Borusyak, Hull, and Jaravel (2025) and the conditions outlined in Goldsmith-Pinkham, Sorkin, and Swift (2020). Accordingly, the standard DiD assumption applies: in the absence of the crisis, plants in high and low-INPRES districts would have followed parallel trends, or exhibited similar elasticities to the shock. I provide supporting evidence for this assumption in the results section below.

5 Results

5.1 Labor Productivity and Output

Table 2 shows results of the first-stage regression of my Bartik instrument from equation (2) on the main variable of interest, $(\text{Post } 1997 \times \text{INPRES})_{dt}$, from the DiD specification in equation (1). The coefficient on the instrument is highly significant, indicating a strong first stage and confirming its relevance. Given that the district INPRES shares are a function of the district-level INPRES intensity and that the cohort shifts are national and common to all districts, the exclusion restriction is plausibly satisfied in this setting.

Tables 3 and 4 show the main results. The first two columns report estimates from the baseline DiD

Table 2: First-Stage Regression Estimating Post 1997 INPRES Exposure Using the Bartik Instrument

	(1)
	Post 1997 × INPRES
Z	1,519***
	(0.995)
Observations	161,547

Bartik IV (Z) = Post 1997 × district share of INPRES schools × national share of fully treated cohorts (born ≥ 1968) in the working-age population (18-60).
Standard errors are clustered at the district level.
*** p<0.01; ** p<0.05; * p<0.10

specification, using district and plant-specific fixed effects, respectively. The last two columns instrument the interaction term using the Bartik instrument from equation (2), again with district and plant-specific fixed effects, respectively. Across all specifications, the coefficient on the interaction between post 1997 and INPRES intensity is positive and statistically significant, indicating that plants in districts with greater INPRES exposure experienced smaller declines, or even gains, in real labor productivity and output during the crisis, relative to plants in districts with lower exposure.

The 2SLS coefficients are only slightly larger than the OLS estimates. The higher magnitudes in the 2SLS results reflect the fact that the effective treatment dose increases over time, as a growing share of the workforce consists of individuals exposed to INPRES schools. This is more clearly evident in Table A.1 in the Appendix, where I re-estimate the OLS model with additional interactions using the national cohort shift variable. The coefficient on the triple interaction (post 1997 × INPRES × shift) is positive and significant across outcomes, indicating that the post 1997 payoff to INPRES rises with the national cohort shifts.

The more conservative within-plant estimates from Model 4 in the two tables indicate that each additional INPRES school in a plant's district led to a 2.8 percent increase in labor productivity (or real output per worker) and a 3.5 percent increase in real output after the 1997 crisis, on average. The fact that both

Table 3: Impact of INPRES on Labor Productivity During the 1997 Crisis

	(1) OLS	(2) OLS	(3) 2SLS	(4) 2SLS
Post 1997 × INPRES	0.0482***	0.0270***	0.0494***	0.0282***
	(0.0149)	(0.00978)	(0.0152)	(0.0100)
Observations	161,547	153,559	161,547	153,559
Year FE	Yes	Yes	Yes	Yes
Sector FE	Yes	Yes	Yes	Yes
District FE	Yes		Yes	
Plant FE	No	Yes	No	Yes

Standard errors are clustered at the district level. Dependent variable is log real output per worker.
2SLS columns instrument post 1997 × INPRES intensity with the Bartik term post 1997 × district share of INPRES schools × national share of fully treated cohorts (born $\geq$ 1968) in the working-age population (18-60).
*** p<0.01; ** p<0.05; * p<0.10

Table 4: Impact of INPRES on Output During the 1997 Crisis

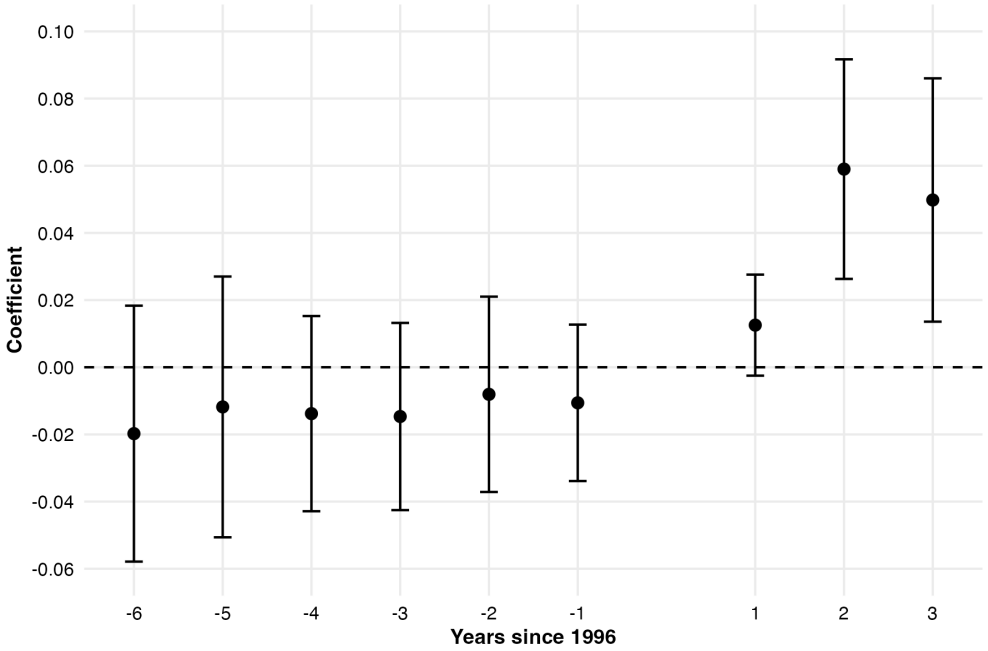
	(1) OLS	(2) OLS	(3) 2SLS	(4) 2SLS
Post 1997 × INPRES	0.0715***	0.0333**	0.0727***	0.0345**
	(0.0251)	(0.0136)	(0.0255)	(0.0139)
Observations	161,547	153,559	161,547	153,559
Year FE	Yes	Yes	Yes	Yes
Sector FE	Yes	Yes	Yes	Yes
District FE	Yes		Yes	
Plant FE	No	Yes	No	Yes

Standard errors are clustered at the district level. Dependent variable is log real output.
2SLS columns instrument post 1997 × INPRES intensity with the Bartik term post 1997 × district share of INPRES schools × national share of fully treated cohorts (born $\geq$ 1968) in the working-age population (18-60).
*** p<0.01; ** p<0.05; * p<0.10

output per worker and total output increased suggests that the gains were not driven by employment cuts. Instead, plants became more productive while maintaining or even expanding their workforce, consistent with the idea that human capital gains from INPRES became more salient and contributed to plant-level resilience during the crisis. I explore this further in the mechanism section below.

Figure 2 shows an event study based on Model 3, with labor productivity as the outcome. The coefficients on the leads of the treatment are small and statistically insignificant, indicating that there were no differential pre-trends between high and low-INPRES districts prior to the 1997 crisis. Results from Model 4 show a similar pattern. This supports the parallel trends assumption and lends credence to the salience hypothesis: INPRES-driven human capital gains did not affect relative performance prior to the crisis but became relevant once the crisis hit. The effects also do not appear immediately in 1997 but become more pronounced in 1998, likely reflecting the lag in plants' ability to implement protective or adaptive responses.

Figure 2: Event-Study of INPRES Effects on Labor Productivity, 1990-1999



Notes: Points are 2SLS estimates of the coefficient on INPRES intensity interacted with each year dummy; 1996 is the omitted year. The interaction term is instrumented with the Bartik term post × district share of INPRES schools × national share of fully treated cohorts (born $\geq$ 1968) in the working-age population (18-60). Vertical bars show 95% confidence intervals with standard errors clustered at the district level. Regression includes district, sector, and year fixed effects.

5.2 Survival

Tables 3 and 4 show results from all plants in the SI sample, excluding those few that appear to change location districts at least once in the data. Models 2 and 4 include plant fixed effects and therefore capture

within-plant impacts, conditional on plants being observed at least once before and after 1997. In other words, Models 2 and 4 estimate the impact on labor productivity and output for surviving plants. Models 1 and 3, which include only district fixed effects, do not condition on survival and instead capture the average effect across all plants, including those that entered or exited the sample.

In other words, Models 2 and 4 capture intensive margin effects, showing how surviving plants became more resilient due to INPRES. Models 1 and 3, by contrast, produce larger coefficients, likely because they reflect both within-plant effects and composition effects, such as more productive entrants replacing less productive exiters. Hallward-Driemeier and Rijkers (2013) found that 1997 the crisis disrupted the usual positive relationship between productivity and survival, leading some productive plants to exit and relatively more productive plants to enter. Consistent with this, regressing an indicator for plant survival, defined as being observed in the SI census, on INPRES exposure yields no significant relationship.

Table 5: Impact of INPRES on Plant Survival During the 1997 Crisis

	(1)	(2)	(3)
	Exit_t	Exit_t	Exit_t
Post 1997 × INPRES		-0.00100	-0.00268
		(0.000905)	(0.00455)
$\log(\text{TFP})_{t-1}$	-0.00110**	-0.00109**	7.50e-05
	(0.000514)	(0.000514)	(0.000386)
Post 1997 × $\log(\text{TFP})_{t-1}$	0.000817**	0.000807**	0.000334
	(0.000362)	(0.000361)	(0.000599)
INPRES × $\log(\text{TFP})_{t-1}$			-0.000693*
			(0.000388)
Post 1997 × INPRES × $\log(\text{TFP})_{t-1}$			0.000179
			(0.000444)
Observations	81,084	81,084	81,084
Year FE	Yes	Yes	Yes
Sector FE	Yes	Yes	Yes
District FE	Yes	Yes	Yes

Standard errors are clustered at the district level. Dependent variable is exit.
*** p<0.01; ** p<0.05; * p<0.10

[See Introduction for discussion of Table 5; further discussion TK]

5.3 Robustness

Because INPRES targeted areas with lower school enrollment that were typically poorer and more rural, and therefore likely to have weaker access to credit, infrastructure, and technology, it is important to verify that the results are not simply capturing pre-existing differences that predate the program. Table A.2 in Appendix A shows that the main crisis estimates remain robust when controlling for district-level pre-crisis measures of capital intensity and financial dependence, including the capital-labor, debt-to-assets, debt-to-output, and interest-to-output ratios. These controls are constructed as district averages rather than plant-level variables so that they capture differential exposure to the crisis through local conditions without conditioning on plant outcomes that INPRES itself may have affected.

Before testing robustness to other potential confounders, I first test whether the main results are driven by a few outlier plants. Table A.4 in Appendix A reports results for labor productivity using a balanced panel of 4,895 plants observed in every year from 1990 to 1999. These estimates are again conditional on survival and are slightly larger than those from the plant fixed effects specification in the full sample. To ensure that these results are not driven by outliers, Table A.5 shows estimates from the full sample after excluding plants with labor productivity below the 5th or above the 95th percentile of the 1996 pre-crisis distribution. The results remain very similar, and are also robust to including sector-by-year fixed effects, suggesting that the main findings are not driven by a few outlier plants or sector-specific shocks.

Finally, I use plant-level deflators instead of a national GDP deflator to account for sectoral, regional, and local price differences during the crisis, following Cali et al. (2024, 2022). These are constructed using product-level price and quantity data from 1994 onward. The results, shown in Table A.6, remain robust, with only a slight reduction in magnitude. This suggests that the estimated effects reflect real differences in plant performance rather than being driven by aggregate price movements. Details on the construction of

these deflators are provided in Appendix B.

[See Introduction for discussion of Table A.3 and other checks; further discussion TK]

Another concern is that my identification strategy implicitly assumes that the district where a plant is located is also the district where its workers were educated. This assumption may be weakened by cross-district migration following the implementation of INPRES. Such migration is common, as individuals often receive their education in the district where their parents reside and later move, typically to urban areas, in search of better job opportunities. Hsiao (2024) shows that these migration patterns contributed to wage gains associated with INPRES. Moreover, individuals with higher levels of education or human capital may be especially likely to migrate, as they can expect higher returns to their skills in more developed districts.

To control for this, I re-run the main regressions with an additional interaction between the post 1997 period and district-level in-migration rates of individuals aged exactly 18-22, using data from the earliest pre-crisis census in 1990. Like in the earlier subsection, I focus on this age group to capture migration patterns specifically for those who would have fully benefited from INPRES. In the 1990 census, the treated cohort (those fully exposed to INPRES during their schooling years) would have been exactly 18-22 years old. This baseline in-migration rate reflects the extent to which INPRES-educated individuals had already moved across districts before the crisis.

Table 6 shows that the results remain robust when controlling for these in-migration rates. In fact, the coefficients in the plant fixed effects specifications increase significantly, possibly because low-INPRES districts attracted more incoming workers educated elsewhere, which inflated their average productivity and biased the original estimates downward. This suggests that the results are not driven by pre-crisis sorting of INPRES-educated workers across districts.

5.4 Sector Import and Export Intensity

[See Introduction for discussion of Table 7; further discussion TK]

Table 6: Impact of INPRES on Labor Productivity Controlling for Pre-Crisis In-Migration Rates

	(1) OLS	(2) OLS	(3) 2SLS	(4) 2SLS
Post 1997 × INPRES	0.0445**	0.0291**	0.0459***	0.0306**
	(0.0172)	(0.0123)	(0.0175)	(0.0125)
Post 1997 × In-migration rate	-0.189	-0.0105	-0.187	-0.00762
	(0.213)	(0.152)	(0.212)	(0.151)
Observations	160,683	152,735	160,683	152,735
Year FE	Yes	Yes	Yes	Yes
Sector FE	Yes	Yes	Yes	Yes
District FE	Yes		Yes	
Plant FE	No	Yes	No	Yes

Standard errors are clustered at the district level. Dependent variable is log real output per worker.
Additional covariate is post 1997 x the five-year in-migration share of 18-22 year olds in each district (1990 census); it controls for differential pre-crisis mobility of the fully INPRES-treated cohorts (born $\geq$ 1968).
2SLS columns instrument post 1997 × INPRES intensity with the Bartik term post 1997 × district share of INPRES schools × national share of fully treated cohorts (born $\geq$ 1968) in the working-age population (18-60).
*** p<0.01; ** p<0.05; * p<0.10

5.5 Pre-Crisis Shift toward Skilled Production

A key concern is that the results might simply reflect pre-INPRES schooling levels rather than the additional education created by INPRES. The program targeted districts with low enrollment, but those areas could still lag in average schooling by the mid-1990s. If low-INPRES districts were already relatively well educated, they would have had a similar human-capital buffer and should have been just as resilient, undermining the interpretation that INPRES-driven gains explain the effect. This view assumes the resilience effect operated purely through primary education gains, rather than through other channels, a point I revisit later.

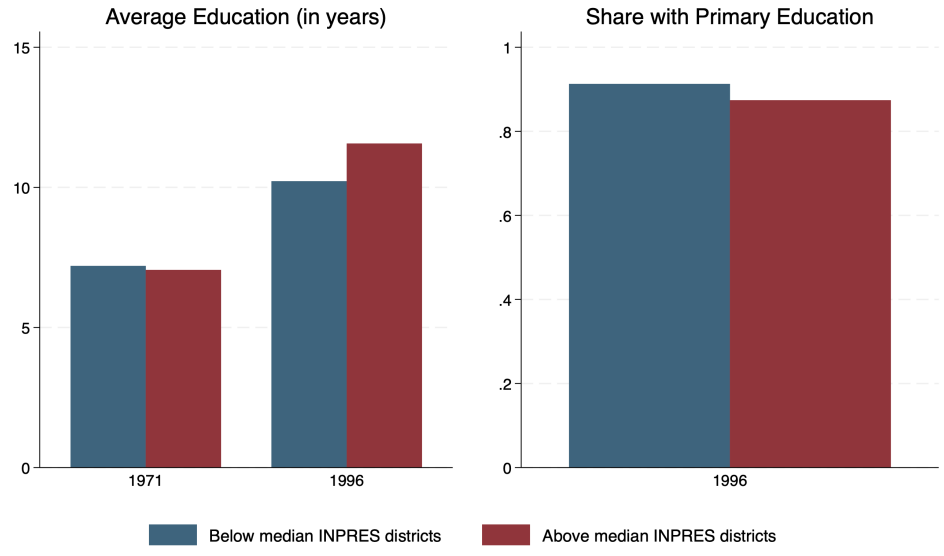
Figure 3 shows that in 1996, workers in above-median INPRES districts still slightly lagged behind those in below-median districts in the share with primary education or higher, although overall rates were high across both groups. In contrast, average years of education were higher in above-median districts, reversing the 1971 pattern (which reflects all adults, not just workers), when below-median districts did slightly better.

Table 7

	(1) Import Intensity (Sector)	(2) Export Intensity (Sector)	(3) Exporter (Plant)
Post 1997 × INPRES	0.0127	0.0244**	0.0291***
	(0.0106)	(0.0106)	(0.0105)
INPRES × Trade orientation	-0.0285	0.0268	
	(0.146)	(0.0633)	
Post 1997 × Trade orientation	-0.357**	0.204**	0.217***
	(0.177)	(0.101)	(0.0318)
Post 1997 × INPRES × Trade orientation	0.279**	0.000967	-0.0320**
	(0.123)	(0.0408)	(0.0151)
Observations	149,808	149,844	125,871
Year FE	Yes	Yes	Yes
Sector FE	Yes	Yes	Yes
Plant FE	Yes	Yes	Yes

Standard errors are clustered at the district level. Dependent variable is log real output per worker.
*** p<0.01; ** p<0.05; * p<0.10

Figure 3: Pre-INPRES and Pre-Crisis Education Levels



Notes: The 1971 data are from the Population Census (for working adults) and sourced from Duflo (2001). The 1996 values are from the SI census and are for plant workers.

These patterns suggest that differences in basic educational attainment are unlikely to fully explain the INPRES effects. Primary education levels were already quite high by the crisis years. If incremental gains in primary schooling were the only channel through which INPRES operated, there would be little variation across plants along this margin. To test this directly, Table 8 re-estimates the main regressions including an interaction between post 1997 and the 1996 share of workers with primary education.

Table 8: Impact of INPRES on Labor Productivity Controlling for 1996 Pre-Crisis District Primary Education Levels

	(1) OLS	(2) OLS	(3) 2SLS	(4) 2SLS
Post 1997 × INPRES	0.0515***	0.0277**	0.0528***	0.0291**
	(0.0170)	(0.0111)	(0.0173)	(0.0114)
Post 1997 × Primary education share (1996)	0.102	0.0189	0.107	0.0238
	(0.172)	(0.115)	(0.172)	(0.116)
Observations	161,399	153,460	161,399	153,460
Year FE	Yes	Yes	Yes	Yes
Sector FE	Yes	Yes	Yes	Yes
District FE	Yes		Yes	
Plant FE	No	Yes	No	Yes

Standard errors are clustered at the district level. Dependent variable is log real output per worker.
2SLS columns instrument post 1997 × INPRES intensity with the Bartik term post 1997 × district share of INPRES schools × national share of fully treated cohorts (born $\geq$ 1968) in the working-age population (18-60).
*** p<0.01; ** p<0.05; * p<0.10

The results remain similar to the original estimates in Table 3, suggesting that the observed resilience effect is not solely explained by pre-crisis primary education levels. This raises the possibility that INPRES may have created broader forms of human capital development beyond simply raising primary schooling attainment. To explore this further, Table 9 shows results from a difference-in-differences analysis using data from the 1990 Population Census and the 1995 Intercensal Population Survey (SUPAS) for cohorts that were fully exposed to INPRES. In 1990, these cohorts were ages 18-22, and by 1995 the same cohorts had reached ages 23-27. The dependent variable is an indicator for employment in skilled production occupations, defined as ISCO-68 major groups 7-8 (craft and related trades; plant and machine operators and assemblers).

Table 9: Pre-Crisis Impact of INPRES on Skilled Production Employment Among Treated Cohorts, 1990-1995

	(1) Skilled prod. occupation	(2) Skilled prod. occupation
Post 1995 × INPRES	0.0456***	0.0455***
	(0.0109)	(0.0109)
Years of schooling		-0.000413
		(0.000302)
Observations	95,114	95,114
Year FE	Yes	Yes
District FE	Yes	Yes
Individual-level controls	Yes	Yes

Dependent variable is an indicator equal to 1 if the individual is employed in a skilled production occupation. Skilled production occupations are defined as ISCO-68 major groups 7-8 (craft and related trades workers; plant and machine operators and assemblers).
Estimates use data from the 1990 Population Census and the 1995 Intercensal Population Survey (SUPAS) for INPRES-exposed cohorts (born 1968-1972; ages 18-22 in 1990 and 23-27 in 1995), restricted to non-movers over the previous five years.
Individual-level controls include age, sex, and an indicator for urban residence.
Standard errors are clustered at the district level.
*** p<0.01; ** p<0.05; * p<0.10

INPRES expanded the local stock of higher-quality human capital in districts where more primary schools were built. This expansion likely occurred not only through increased primary school enrollment, but also through broader improvements in human capital formation. For instance, newer schools may have offered better infrastructure or instruction, and early access to schooling may have raised aspirations, encouraging further vocational preparation. By the 1990s, a growing share of the workforce consisted of younger cohorts who had attended INPRES schools, increasing the relevance of these gains during the crisis.

As shown in Table 9 and discussed in Section 5.2, INPRES appears to have contributed to a shift toward skilled production occupations before the crisis, and this channel, rather than differences in primary education levels, may have driven the resilience effect during the downturn. Table A.7 in the Appendix further shows that in 1996, plants in high-INPRES districts were more likely to have workers engaged in training, consistent with broader improvements in workforce skills beyond primary education. Taken together, these findings suggest that INPRES not only expanded basic schooling but also facilitated movement into more

skill-intensive forms of work.

5.6 Expanding Employment Amid Crisis

To examine whether this expansion of human capital affected how plants adjusted during the crisis, I first test whether INPRES-exposed plants expanded employment more or contracted less than others. Table 10 re-estimates the main specifications with the change in plant workforce size (in logs) as the dependent variable using the balanced panel. The interaction between post 1997 and INPRES intensity is positive and significant across all four specifications. The results suggest that one additional INPRES school per 1,000 children increased a plant's employment growth by about 1 percent after the crisis.

Table 10: Impact of INPRES on Change in Plant Workforce During the 1997 Crisis

	(1) OLS	(2) OLS	(3) 2SLS	(4) 2SLS
Post 1997 × INPRES	0.00920**	0.00984***	0.00886**	0.00951***
	(0.00356)	(0.00352)	(0.00351)	(0.00347)
Observations	44,048	44,048	44,048	44,048
Year FE	Yes	Yes	Yes	Yes
Sector FE	Yes	Yes	Yes	Yes
District FE	Yes		Yes	
Plant FE	No	Yes	No	Yes

Standard errors are clustered at the district level. Dependent variable is annual log change in total workforce.
Sample is limited to a balanced panel of 4,895 plants observed in every year from 1990 to 1999.
2SLS columns instrument post 1997 × INPRES intensity with the Bartik term post 1997 × district share of INPRES schools × national share of fully treated cohorts (born $\geq$ 1968) in the working-age population (18-60).
*** p<0.01; ** p<0.05; * p<0.10

As shown in the results section, high-INPRES plants experienced real relative productivity gains during the crisis that were not driven by reductions in employment. Given that the crisis was a negative demand shock, plants could have responded by downsizing, especially since higher labor productivity would have allowed them to maintain output with fewer workers. Instead, Table 10 shows that INPRES-exposed plants

expanded their workforce relative to others. A possible explanation is labor hoarding, possibly reflecting an effort to retain more productive workers in anticipation that the shock would be temporary.

5.7 Retention of Educated Workers

Beyond expanding employment overall, INPRES-exposed plants also appear to have retained a more educated workforce during the crisis. I examine this by re-estimating the main regression specifications using the number of primary-educated workers as the dependent variable. As shown in Table 11, high-INPRES plants employed significantly more primary-educated workers after the crisis.

Table 11: Impact of INPRES on Number of Primary Educated Workers During the 1997 Crisis

	(1)	(2)
	OLS	OLS
Post 1997 × INPRES	0.0291**	0.0193***
	(0.0115)	(0.00725)
Observations	51,257	47,449
Year FE	Yes	Yes
Sector FE	Yes	Yes
District FE	Yes	
Plant FE	No	Yes

Standard errors are clustered at the district level. Dependent variable is number of workers with primary education or above (in logs).
Regressions are estimated only using 1995-1997 data, when worker education is observed in the SI census.
*** p<0.01; ** p<0.05; * p<0.10

These results suggest that plants recognized the value of their more skilled workforce and chose to retain them, which likely contributed to their ability to weather the downturn more effectively. This behavior is consistent with labor hoarding, where firms employing higher-skilled workers are often more reluctant to lay them off during crises, particularly when facing high adjustment costs (Guimbert and Oostendorp, 2016). This pattern was also observed among Southeast Asian firms during the 1997 crisis (Hallward-Driemeier, Dwor-Frécaut, and Colaco, 2000).

5.8 Falling Wages for Production Workers

Next, I examine how the crisis affected the relative price of different types of work, comparing nonproduction roles (managers, technicians, clerical staff) to production roles (machine operators, assembly line workers). I re-estimate the main specification using the difference between the log average wages of nonproduction and production workers as the dependent variable. Table 12 shows the results. The coefficient on the main interaction term is positive and statistically significant, indicating that each additional INPRES school per 1,000 children increased the wage gap between nonproduction and production workers by about 4 percent among surviving plants.

Table 12: Impact of INPRES on the Wage Gap Between Nonproduction and Production Workers During the 1997 Crisis

	(1) OLS	(2) OLS	(3) 2SLS	(4) 2SLS
Post 1997 × INPRES	0.0547***	0.0392**	0.0549***	0.0396**
	(0.0181)	(0.0173)	(0.0180)	(0.0173)
Observations	115,509	107,944	115,509	107,944
Year FE	Yes	Yes	Yes	Yes
Sector FE	Yes	Yes	Yes	Yes
District FE	Yes		Yes	
Plant FE	No	Yes	No	Yes

Standard errors are clustered at the district level. Dependent variable is the difference between the log average wages of nonproduction and production workers.
2SLS columns instrument post 1997 × INPRES intensity with the Bartik term post 1997 × district share of INPRES schools × national share of fully treated cohorts (born $\geq$ 1968) in the working-age population (18-60).
*** p<0.01; ** p<0.05; * p<0.10

This suggests that the relative value of production work, which had increased before the crisis in high-INPRES districts, declined compared to supervisory or nonproduction roles during the crisis. Although wages fell across all worker types, Table 13 tests whether INPRES affected how sharply they declined. The results show that high-INPRES plants experienced steeper overall wage declines, driven primarily by falling wages among production workers. In contrast, wages for nonproduction workers were not differentially

affected.

Table 13: Impact of INPRES on Wage Changes by Worker Type During the 1997 Crisis

	(1) All	(2) Prod	(3) Nonprod
Post 1997 × INPRES	-0.0225**	-0.0283**	-0.00585
	(0.0100)	(0.0123)	(0.0130)
Observations	28,310	28,310	28,310
Year FE	Yes	Yes	Yes
Sector FE	Yes	Yes	Yes
District FE	Yes	Yes	Yes

Standard errors are clustered at the district level. Dependent variable is the log change in average wages for each worker type (all, production, and nonproduction).
Sample is restricted to a balanced panel of 4,895 plants observed in every year from 1990 to 1999, and further limited to those plants for which all three wage types are available.
*** p<0.01; ** p<0.05; * p<0.10

5.9 Putting It All Together: A Stylized Model of Firm Adjustment

Summarizing the INPRES-induced changes before and during the crisis, three main patterns emerge:

(i) In the years leading up to the crisis, INPRES-exposed cohorts were more likely to enter skilled production occupations, and plants in high-INPRES districts were more likely to have workers engaged in training.

(ii) High-INPRES plants retained or expanded their workforce more than low-INPRES plants, and the workers they retained were more likely to have completed primary education.

(iii) Although wages declined across the board, the gap between nonproduction and production wages widened, driven primarily by a sharper decline in production wages.

Given the pre-crisis increase in INPRES-exposed cohorts working in skill-intensive production work, the relative decline in production wages likely reflects the greater local supply of such workers in high-INPRES

districts during the crisis. This abundant supply helps explain both the relative expansion of workforces by plants in these districts and the retention of primary-educated workers.

Building on these patterns, I outline a simple model to explain how this contributed to crisis resilience, focusing on production workers as the sole labor input. Suppose the local economy has two types of production workers, W_H and W_L (high and low human capital, respectively), and that the supply of W_H increased by a factor $\Delta W_H > 0$. Plants are assumed to choose the number of high and low human capital production workers each period to maximize profits, with output given by the production function $Y = A_t F(\phi_t W_H, W_L, \bar{K})$, where A_t is a demand shifter and ϕ_t is a skill-biased term that captures human capital salience. I assume $\phi_t = 1$ prior to the crisis and that ϕ_t rises during the crisis.

When the financial crisis hits, A_t falls, which decreases the marginal product of labor for both types of workers. Under the human capital salience hypothesis, however, ϕ_t rises during the crisis, so W_H becomes relatively more valuable, since higher skills complement the capital and organizational adjustments plants undertake in response to the shock. As a result, the marginal product of W_L falls more than that of W_H.

High-INPRES plants have access to an additional supply of high human capital labor, $\Delta W_H > 0$, which keeps wages for W_H relatively low despite their higher marginal product during the crisis. This abundance allowed plants to retain more skilled production workers and sustain output in the face of the negative shock. In this way, the abundance of skilled production labor, combined with rising returns to those skills and lower relative wages for W_H, enhanced the resilience of high-INPRES plants during the crisis.

This model is loosely based on the framework developed by Kremer and Maskin (1996), who show that when workers of different skill levels are imperfect substitutes, and tasks within firms are complementary and vary in their sensitivity to skills, an increase in the average skill level can lead to greater wage inequality.

5.10 Salience of Human Capital During the Crisis

Finally, I provide additional supporting evidence for the human capital salience hypothesis, the idea that skills become more important during a crisis, which was taken as given in the model above. The widening wage gap between production and nonproduction workers during the crisis is consistent with this mechanism, but that comparison is limited to two broad worker categories. A more direct test uses observed primary education levels, even though these are not the specific channel through which the INPRES effect operates. By the mid-1990s, average primary schooling rates were high, but there was still some variation across plants (see Figure A.2 in Appendix A), allowing the salience hypothesis to be examined more directly.

Table 14 shows this test. Using the only three years in which the SI census records worker schooling (1995-1997), I regress labor productivity on the share of workers with at least primary education, estimating separate coefficients for the pre-crisis years (1995-1996) and the crisis year (1997). Because data are collected at year-end, the 1997 coefficient captures plants operating under crisis conditions. Education is positively correlated with productivity in both periods, but the coefficient in 1997 is statistically significantly larger than in the previous period.

This widening premium suggests that the returns to a more educated workforce increased during the crisis. Of course, I cannot include plant fixed effects in these regressions, so the estimates compare different plants across the pre-crisis and crisis periods, and the sample size is larger in the pre-crisis years. Firm entry and exit could bias these estimates, but the relationship between education and productivity remains clearly positive and appears to strengthen during the crisis.

Table 14: Relationship between Plant Workforce Education & Labor Productivity, Pre vs. During Crisis

	(1)	(2)
	1995-96	1997
Share of workers with primary educ. & above	0.587***	0.729***
	(0.0761)	(0.0834)
Observations	34,641	16,714
Year FE	Yes	
Sector FE	Yes	Yes
District FE	Yes	Yes

Standard errors are clustered at the district level. Dependent variable is log real output per worker.
Coefficient on share of workers with primary education is statistically significantly higher in 1997 compared to 1995-96.
*** p<0.01; ** p<0.05; * p<0.10

A Appendix: Figures and Tables

Figure A.1: Regional Distribution of Manufacturing Plants and INPRES Intensity

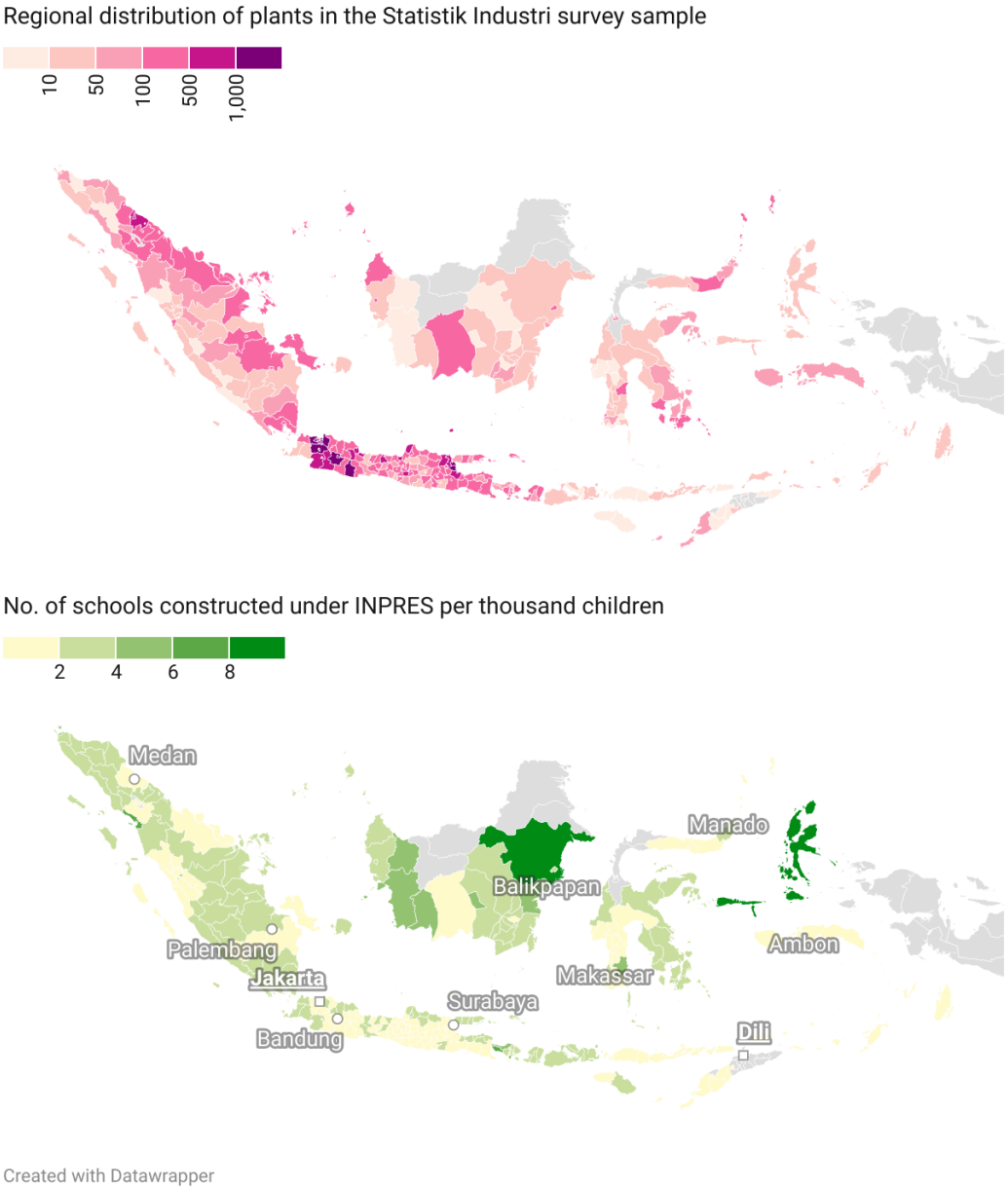


Table A.1: Resilience Effect of INPRES Scales with National Share of Treated Working-Age Cohorts

	(1) Labor Prod.	(2) Labor Prod.	(3) Output	(4) Output
Post 1997 × INPRES × Shift	0.748**	0.664**	0.883*	0.875***
	(0.364)	(0.307)	(0.455)	(0.335)
Observations	161,547	153,559	161,547	153,559
Year FE	Yes	Yes	Yes	Yes
Sector FE	Yes	Yes	Yes	Yes
District FE	Yes		Yes	
Plant FE	No	Yes	No	Yes

Standard errors are clustered at the district level. Dependent variables are log real output per worker and log real output.
Shift is the national share of fully treated cohorts (born $\geq$ 1968) in the working-age population (18-60). This is the same *shift* term used in the Bartik instrument in the 2SLS results.
*** p<0.01; ** p<0.05; * p<0.10

Table A.2: Impact of INPRES on Labor Productivity Controlling for Pre-Crisis Capital Intensity and Financial Dependence

	(1) log(K/L)	(2) Debt/assets	(3) Debt/output	(4) Interest/output	(5) All controls
Post 1997 × INPRES	0.0297***	0.0317***	0.0272***	0.0271***	0.0314***
	(0.0109)	(0.0112)	(0.00996)	(0.00981)	(0.0115)
Post 1997 × Control	0.0123	0.0884	2.22e-05	0.186	
	(0.0219)	(0.0956)	(0.000168)	(0.396)	
Observations	153,437	153,069	153,454	153,460	153,019
Year FE	Yes	Yes	Yes	Yes	Yes
Sector FE	Yes	Yes	Yes	Yes	Yes
Plant FE	Yes	Yes	Yes	Yes	Yes

Standard errors are clustered at the district level. Dependent variable is log real output per worker.
Columns (1)-(4) each control for a district-level average of a pre-crisis baseline variable interacted with post 1997: (1) capital-labor ratio (in logs) [1995], (2) debt-to-assets ratio [1996], (3) debt-to-output ratio [1996], and (4) interest-to-output ratio [1996]. Column (5) includes all four interacted baseline variables jointly.
*** p<0.01; ** p<0.05; * p<0.10

Table A.3

	(1) Foreign ownership	(2) Govt. ownership	(3) Group membership
Post 1997 × INPRES	0.0302**	0.0304**	0.0278**
	(0.0121)	(0.0123)	(0.0120)
Control (baseline)	1.499***	0.253***	0.709***
	(0.151)	(0.0954)	(0.0895)
INPRES × Control	-0.207**	0.0616	-0.136***
	(0.0899)	(0.0505)	(0.0407)
Post 1997 × Control	0.139	-0.0701	0.0664
	(0.156)	(0.0987)	(0.0664)
Post 1997 × INPRES × Control	0.0229	0.0228	0.0128
	(0.0844)	(0.0508)	(0.0361)
Observations	126,410	126,410	117,243
Year FE	Yes	Yes	Yes
Sector FE	Yes	Yes	Yes
District FE	Yes	Yes	Yes

Standard errors are clustered at the district level. Dependent variable is log real output per worker.
*** p<0.01; ** p<0.05; * p<0.10

Table A.4: Impact of INPRES on Labor Productivity Using Balanced Plant Panel, 1990-1999

	(1) OLS	(2) OLS	(3) 2SLS	(4) 2SLS
Post 1997 × INPRES	0.0387***	0.0387***	0.0392***	0.0394***
	(0.0119)	(0.0117)	(0.0122)	(0.0120)
Observations	48,944	48,944	48,944	48,944
Year FE	Yes	Yes	Yes	Yes
Sector FE	Yes	Yes	Yes	Yes
District FE	Yes		Yes	
Plant FE	No	Yes	No	Yes

Standard errors are clustered at the district level. Dependent variable is log real output per worker.
Sample is limited to a balanced panel of 4,895 plants observed in every year from 1990 to 1999.
2SLS columns instrument post 1997 × INPRES intensity with the Bartik term post 1997 × district share of INPRES schools × national share of fully treated cohorts (born ≥ 1968) in the working-age population (18-60).
*** p<0.01; ** p<0.05; * p<0.10

Table A.5: Impact of INPRES on Labor Productivity Using Trimmed Sample

	(1) OLS	(2) OLS	(3) 2SLS	(4) 2SLS
Post 1997 × INPRES	0.0470***	0.0265***	0.0484***	0.0278***
	(0.0131)	(0.00883)	(0.0135)	(0.00916)
Observations	143,692	135,754	143,692	135,754
Year FE	Yes	Yes	Yes	Yes
Sector FE	Yes	Yes	Yes	Yes
District FE	Yes		Yes	
Plant FE	No	Yes	No	Yes

Standard errors are clustered at the district level. Dependent variable is log real output per worker. Excludes all plants with labor productivity below the 5th or above the 95th percentile of the 1996 pre-crisis distribution.
2SLS columns instrument post 1997 × INPRES intensity with the Bartik term post 1997 × district share of INPRES schools × national share of fully treated cohorts (born $\geq$ 1968) in the working-age population (18-60).
*** p<0.01; ** p<0.05; * p<0.10

Table A.6: Impact of INPRES on Labor Productivity Using Plant-Level Price Deflators

	(1) OLS	(2) OLS	(3) 2SLS	(4) 2SLS
Post 1997 × INPRES	0.0366***	0.0217**	0.0369***	0.0218**
	(0.0107)	(0.00869)	(0.0110)	(0.00890)
Observations	90,691	86,967	90,691	86,967
Year FE	Yes	Yes	Yes	Yes
Sector FE	Yes	Yes	Yes	Yes
District FE	Yes		Yes	
Plant FE	No	Yes	No	Yes

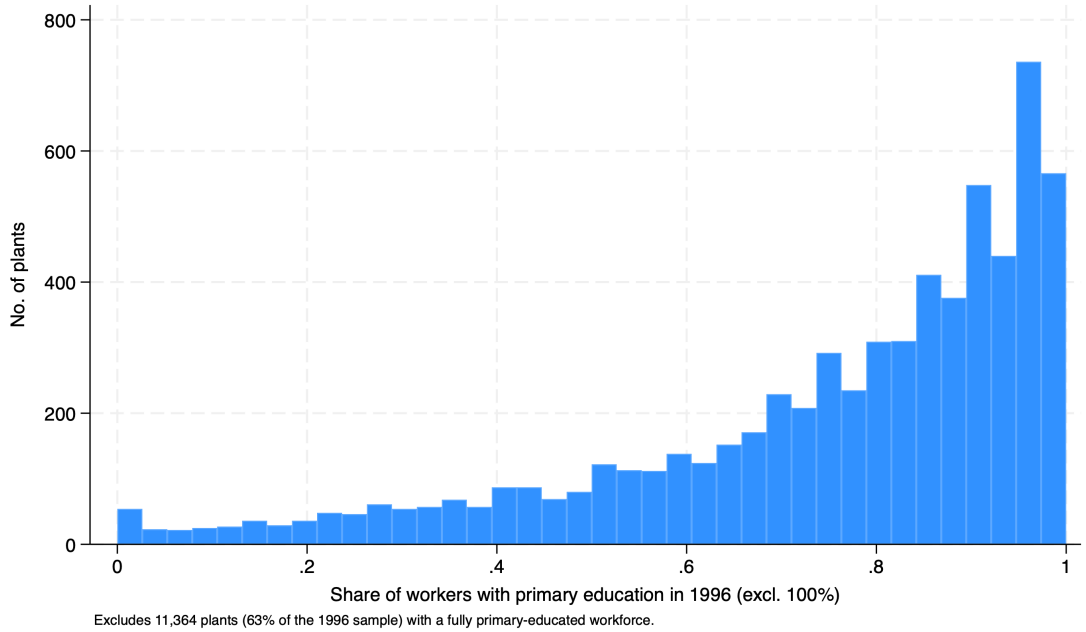
Standard errors are clustered at the district level. Dependent variable is log real output per worker, deflated with plant-specific Tornqvist price indices.
2SLS columns instrument post 1997 × INPRES intensity with the Bartik term post 1997 × district share of INPRES schools × national share of fully treated cohorts (born $\geq$ 1968) in the working-age population (18-60).
*** p<0.01; ** p<0.05; * p<0.10

Table A.7: INPRES and Pre-Crisis Plant Training in 1996

	(1) Training	(2) Training
INPRES	0.0274**	0.0353***
	(0.0112)	(0.0106)
Workers		0.123***
		(0.00612)
Labor Prod.		0.0503***
		(0.00763)
Observations	16,378	16,378
Sector FE	Yes	Yes

Standard errors are clustered at the district level. Dependent variable is an indicator equal to 1 if the plant provided any worker training in 1996.
Column (2) controls for log total workers and log real output per worker in 1996.
*** p<0.01; ** p<0.05; * p<0.10

Figure A.2: Distribution of Primary Schooling Among Workers in 1996 (Excl. 100%)



B Appendix: Construction of Plant-Level Price Deflators

To account for sectoral, regional, and local price variation during the crisis period, I construct plant-level price deflators using product-level price and quantity data from 1994 onward. The methodology follows Cali et al. (2024, 2022), which builds on Eslava et al. (2004) and Mertens (2019).

The raw dataset requires substantial preprocessing. Units of measurement vary across years for the same products, and in some cases, units are ambiguous or not easily convertible into consistent metrics such as volume, weight, count, area, or length. To ensure consistency, I first retain only products that appear in all years for each plant. I then keep only products with clear and convertible units that can be standardized into one of five categories: volume, weight, count, area, or length. Finally, following Cali et al. (2024), I drop products where the implied unit price changes by more than a factor of 10 between consecutive years.

Using the cleaned dataset, I construct a Tornqvist price index for each plant to capture variation in price changes across its 9-digit products. The plant-level index is defined as:

$$\pi_{it} = \prod_{p=1}^{n} \left(\frac{P_{pit}}{P_{pit-1}} \right)^{0.5(S_{pit}+S_{pit-1})} \pi_{it-1},$$

where P_{pit} is the unit value (total revenue divided by total quantity) of product p in plant i at time t, and S_{pit} is the share of product p in the plant's total product sales at time t. The price index π_{it} evolves as a weighted geometric mean of product-level price growth, with weights equal to the average sales share in periods t and $t-1$.

For missing values, I follow Cali et al. (2024) and impute price growth using the average change observed in the same sector and district. I use these plant-level deflators to adjust outcome variables and re-estimate the main regression specifications. Results using these adjusted outcomes are reported in Table A.6.

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