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Immigration and Local Inflation

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Immigration and Local Inflation
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ABSTRACT: We use a shift-share approach to estimate the impact of inward immigration on local inflation in the United States. We find that a higher rate of immigration reduces inflation, lowering it by about 0.1 to 0.2 percentage points following a doubling of immigration. Higher immigration flows also lower local goods inflation, increase local housing and utilities inflation, and have no statistically significant impact on inflation in other services. Effects are approximately two and three time larger for working age and low-education immigrants. We do not detect a statistically significant impact of more educated immigrants on overall inflation, but they do increase local housing inflation. Our results can be jointly rationalized by a simple general equilibrium model where the substitutability of capital and labor varies across industries but capital is fixed in the short run.

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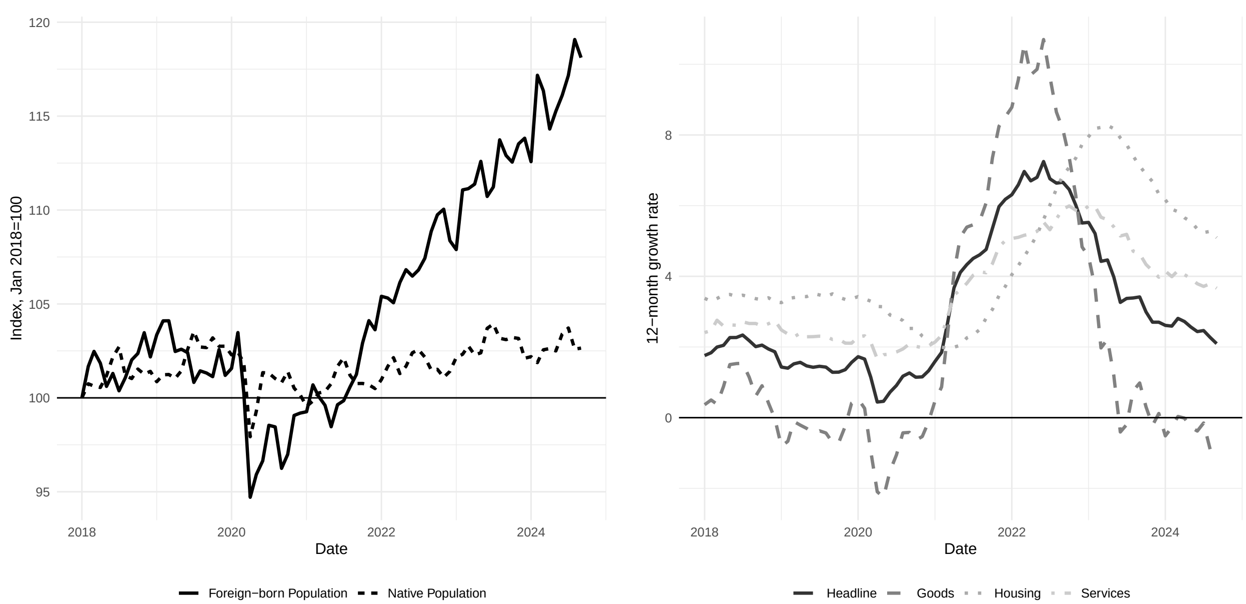
1 Introduction

In recent years, immigration has accelerated significantly from pre-pandemic levels. Since 2021, new immigrants have expanded the labor force by 4.6 million workers – some three percent of the labor force. At the same time, inflation surged, with twelve-month core personal consumption expenditure (PCE) inflation peaking at 5.6 percent in early 2022 before declining to a little over 2 percent today. Relative prices also varied considerably during this time period, with increases in services and housing inflation lagging that of goods (See Figure 1). These correlations raise a natural question: what impact has immigration had on inflation in the US economy?

Figure 1: Immigration and inflation in the US, 2018-2024



(a) Labor force, by nativity. **(b)** PCE inflation, headline and key components

Notes: Figure (a) presents the foreign-born and native population of the United States indexed to Jan 2018. Figure (b) presents the 12-month growth rate in PCE and its components. Data is from the US Census and Bureau of Economic Analysis.

The answer is not immediately obvious since immigration increases both supply and demand. Inflows of workers boost supply, alleviating labor shortages, moderating wage pressures, and easing inflation. But more immigration also increases demand for goods and services, driving up output, wages, and inflation. Moreover, the balance of supply and demand may vary considerably across products, affecting relative prices even more than the overall price level.

This paper attempts to address this question by looking at the relationship between immigration and inflation in 306 metropolitan statistic areas in the United States between 2009 and 2022. The key challenge in interpreting this relationship is that immigration and local inflation are likely jointly determined, with each having a causal impact on the other. We thus use a shift-share approach to isolate plausibly exogenous variation in local immigration.[1] We construct an instrument for local immigration from average national migration by country weighted by the 2005 national origin share for each Metropolitan statistical area (MSA). This instrument will be strong if pre-existing connections between foreign countries and US cities mean that immigrants from some countries are more likely to move to particular cities. The instrument will isolate exogenous variation in immigration if fluctuations in economic conditions of individual US metro areas are not sufficient to impact overall outflows from origin countries.

We find that immigration reduces local inflation relative to other MSAs, but that the effects are moderate and are mainly driven by low-education immigrants. Doubling the rate of immigrant inflows reduces overall inflation by about 0.1 to 0.2 p.p.. When restricted to non-college educated migrants the effects are around three times as big. In contrast, the impact of high-education immigration on overall inflation is statistically insignificant. We also find that immigration lowers local goods inflation, increases housing inflation, and has a negligible effect on services. However, for all products, less educated immigrants tend to be more disinflationary than more educated immigrants. Finally, the effects of immigration on local inflation tend to be transitory, with effects dissipating after one year. One exception is the price of housing, where the (dis)inflationary effects of

[1] First proposed by Bartik (1991), the shift-share approach has been widely used in the literature on immigration, such as in Card (2001), Card (2009), Peri and Sparber (2009), Peri et al. (2015), Monras (2020), Caiumi and Peri (2024), and others.

(low-) high-education immigrants can last several years

To interpret these results together we develop a simple static model where immigration is a positive shock to labor supply, but production is non-homothetic. This can help square our results with standard views on the substitutability of labor and capital in different industries. For example, labor and capital are generally thought of as more substitutable in the production of goods than in services when labor supply increases. And so if capital is fixed in the short term, the short-term elasticity of supply of goods should be larger than that of services, leading to a decline in the relative price of goods versus services. This is exactly what we see in our estimated results. A similar argument applies to the prices of housing and utilities, where labor and capital are poorer substitutes and also see the largest relative price rises. So too for our results on high- versus low-education labor, the latter of which is generally considered to be a better substitute for capital and so less-educated immigrants should lower prices by more than high-education immigrants. This is also what we see in the data. Our results are also consistent with capital being elastically supplied in the long run. Relative prices in all sectors are statistically indistinguishable with their pre-immigration levels three years after the shock. The one exception is housing, consistent with land being a fixed factor even in the long run.

Related Literature. This paper builds on the rich literature exploring the economic implications of immigration. The most prominent line of work in this literature focuses on the impact on the wages and employment of native workers (Ottaviano and Peri, 2012; Manacorda et al., 2012; Borjas, 2003; Card, 1990; Clemens and Hunt, 2019; Borjas, 2017; Borjas, 2017).[2] Several papers in this literature construct "shift-share" instruments similar to ours to compare labor market outcomes across regions, such as in Card (2001), Card (2009), Peri and Sparber (2009), Peri et al. (2015), Monras (2020), and Caiumi and Peri (2024). Overall, although many studies conclude that the impact of immigration

[2]For further details, chapter 5 of National Academies of Sciences, Engineering, and Medicine (2017) is a popular paper summarizing the effect of immigrants on labor market outcomes in the US and covers a wide range of research in this area.

on average wages and employment is small, a high degree of consensus exists that less educated workers are more vulnerable than others to inflows of new immigrants. Particularly, low-education immigration is more likely to substitute and lower the wages of less educated native workers, while there is a greater likelihood of complementarity between high-education immigrant and native workers raising wages. Consistent with the findings from this literature, we find that less educated immigrants tend to be more disinflationary than more educated immigrants.

Work looking specifically at the impact of immigration on inflation is more limited, but has generally also found deflationary impacts of inward immigration. Lach (2007) uses cross-section variation in retail store-product prices and finds that a surge in former Soviet Union immigrants to Israel in 1990 lowered prices, consistent with our results on goods inflation. In a closely related paper, Cortes (2008) uses a an identification strategy similar to ours to study the relationship between the stock of low-skilled immigrants and prices across US cities, finding that immigration lowers the price of immigrant-intensive services. Since only decadal census data were available at the time, she focuses on long-run relationships which, we argue, obscures much of the impact of immigration on prices. Improved data[3] allow us to revisit this question, identifying effects of immigration "flows" at an annual frequency which, for the most comparable groups, we calculate to be approximately twenty times as large. To the best of our knowledge, this paper is the first to address this question for the United States with this degree of temporal and spatial resolution. More recently, Cheremukhin et al. (2024) adopt a complementary approach to this question. Using a quantitative model, they argue that the post-pandemic immigration surge would have have largely offsetting supply and demand effects.

The rest of the paper is organized as follows. Sections 2 and 3 describe the methodology

[3]The American Community Survey, our principal source of data on annual immigration, was not fully implemented until 2005. In contrast, Cortes (2008) had only census records for 1980, 1990, and 2000 available. The higher frequency of observations in the ACS, as well as refinements in geographic coverage, mean that our identifying variation uses around forty times as many observations.

and data. Section 4 presents the results. Section 5 interprets those results through the lens of a simple model. Section 6 concludes.

2 Methodology

We model the effect of immigration on inflation in MSA i and year t as follows:

$$\Delta \log(P_{it}) = \beta \frac{\Delta F_{it}}{Pop_{i,t-1}} + X_{it} + \epsilon_{it} \tag{1}$$

where inflation defined as the log difference in the price level between year t and $t-1$, $\Delta \log(P_{it})$. ΔF_{it} is the change in foreign-born population in MSA i between t and $t-1$ and $Pop_{i,t-1}$ is the total population in MSA i in $t-1$.[4] X_{it} is a set of controls, including time fixed effects in all specifications to control for national inflation variation. MSA fixed effects are accounted for by first-differencing.

The coefficient β measures the impact of net immigration per capita on inflation. However, estimating Equation (1) by ordinary least squares (OLS) risks producing biased estimates. This will arise if unobserved local characteristics which affect inflation (captured in the term ϵ_{it}) are correlated with immigration, and could emerge from a number of sources. Most obviously, there could be direct reverse causality: immigrants may be attracted to locations where inflation is low. Alternatively, other economic factors, such as increased demand for a locally concentrated industry (e.g. information technology in the San Francisco Bay Area), could constitute an omitted variable, both attracting new immigrants and impacting inflation.

To address this problem, we use an instrumental variables strategy that identifies sources of variation from changes in foreign-country immigration which are plausibly uncorrelated with local economic conditions. We do this by constructing an instrument which

[4] The lagged divisor avoids some immediate endogeneity issues.

is the weighted average of immigration to the US from all foreign countries, where the weights are the MSA-level residents from those countries in a pre-sample base year. This sort of method is widely used in the literature on immigration, including in Altonji and Card (1991), Card (2001), and Cortes (2008).

Specifically, let $sh_{c,i,t'}$ be the foreign-born population from country c, living in MSA i as of reference year t', as a share of their total population in the United States in year t':

$$sh_{c,i,t'} = \frac{F_{c,i,t'}}{\sum_i F_{c,i,t'}} \tag{2}$$

where $F_{c,i,t}$ is the foreign-born population from country c in MSA i as of year t. Total national immigration from country c is thus:

$$\Delta F_{c,t} = \sum_i F_{c,i,t} - \sum_i F_{c,i,t-1} \tag{3}$$

We take 2005 as our base year (our sample otherwise runs from 2009 to 2022) and construct our shift-share using $sh_{c,i,2005}$ as the "share" and $\Delta F_{c,t}$ as the “shifts”. That is:

$$instrument_{it} = \frac{\sum_c sh_{c,i,2005} \Delta F_{c,t}}{Pop_{i,t-1}} \tag{4}$$

where the denominator is just a normalizer, consistent with the dependent variable in equation (1).

To be valid, the instrument must be strong and plausibly exogenous. Although strength is ultimately an empirical question, the pre-existing links between immigrant communities in US cities and their origin countries provide some conceptual justification for this approach. Intuitively, if immigrants are more likely to settle where their compatriots have previously done so, a shift in national inflows towards a particular foreign country should map into an increase in relative local foreign-born population in the US cities which have previously hosted migrants from that country. For example, an increase in total immigration to the US from Armenia relative to that from Somalia will likely be correlated with more immigration into Los Angeles (which has a large Armenian diaspora) than Minneapolis (with a relatively large Somali population). The argument for exogeneity relies on US MSAs being sufficiently small that local economic conditions do not drive out-migration from particular countries. One concern is that if immigration is too highly or origin countries are very small, then economic fluctuations in US MSAs could have a meaningful effect on outmigration from origin countries. To address this, later we decompose our estimator into the contributions from different regions, and show that the most important drivers of our results are large countries and with relatively diffuse immigration patterns (e.g. Mexico and India).

3 Data

This section describes the data and the construction of the key variables.

Our inflation data comes from the U.S. Bureau of Economic Analysis's Regional price parities (RPPs). Regional price parities are price indexes that measure geographic price level differences within the United States. These are constructed from local prices, aggregated by local expenditure weights from the BEA's personal consumption expenditure (PCE) series. As such, RPPs can be thought of as the local analogue to the national PCE inflation index.[5] The BEA publishes RPP indices for four disjoint components of the main index: goods, housing, utilities, and other services (i.e. excluding housing and utilities). We conduct our analysis across metropolitan statistical areas (MSAs), for which data is available annually from 2009 onwards.

The RPPs are not without limitations, some of which hinder the direct comparison of the

[5]Technically, the RPPs are strictly *relative* price indices, so to compute local inflation, we multiply the change in RPPs by the change in the U.S. PCE price index, as suggested by the BEA. See BEA, 2023 for details.

RPP to national PCE. Most notably, the housing component for the RPPs only includes tenant's rental costs, not owner-imputed rents. Although recent immigrants themselves may be typically renters, omitting owner-imputed rents will exclude the impact that increased demand for (or supply of) housing has on local home-owners' housing costs. Moreover, the microdata used for the housing component of the RPP is a sample from local areas, adjusted for observable characteristics of the dwelling.[6] If there are changes in the composition of housing which are not captured by these factors which are correlated with inward immigration, then the RPP methodology will reflect those rather than just pure price changes. Moreover, the RPPs exclusive focus on rental prices means that it measures prices with a time lag due to long-term leases/rental contracts. So, we complement our analysis using RPPs with two more detailed measures of housing prices, the Zillow Home Value Index and the Zillow Observed Rent Index (ZHVI and ZORI respectively). These measures uses property-level actual and market-implied prices for over 100 million properties in the United States. As such, they adjust for composition changes at source. The indices use the individual prices to compute the costs of "typical" home value by MSA, i.e. those in the 35th to 65th percentile of prices.[7] The index also reflects current "market" prices, rather than that captured from a mix of long-term leases that have or have not been renewed recently.

Our population and immigration data come from the 2005 to 2022 American Community Surveys (ACS), all obtained through IPUMS (Ruggles, et al. 2015). We define a foreign-born resident as a person born in a country other than the U.S. (excluding outlying U.S. territories). We sum the person weights by MSA and birthplace to compute $F_{c,i,t}$. In some cases the ACS birthplaces are defined by country group (for example, the West Indies). We include in the analysis all MSAs that can be identified in ACS, using the 2013 definitions for metropolitan statistical areas from the U.S. Office of Management and Budget.

[6]Specifically: type of building, number of rooms and bedrooms in the unit, and age of the unit.
[7]Indices of other parts of the distribution are also available, which we use in robustness tests.

We have consistent MSA delineations in the ACS from 2005 onwards, thus we use 2005 as our reference year for the "shift-share" instrument. We consider six subgroups of immigrants in the analysis using the age and educational attainment variables in the ACS: all foreign-born residents; working age foreign-born residents (Ages 18 to 64); working age foreign-born residents with no high school diploma; working age foreign-born residents with a high school diploma or higher; working age foreign-born residents with no college education; working age foreign-born residents with some college education.

Together, our analysis includes 306 MSAs and years 2009 to 2022. Appendix Table A1 presents summary statistics for the data.

4 Results

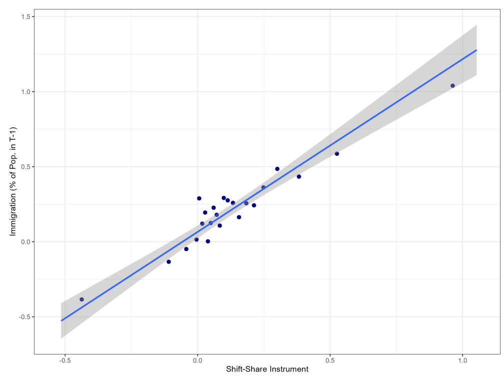
4.1 First Stage

Figure 2 shows the first-stage correlations between the shift-share instrument and net immigration per capita. The positive correlations are clearly visible. Appendix Table A2 reports the first stage estimates. The F statistics for these first-stage partial correlations range from 21 to 43 across immigrant subgroups. The instrument is strong, with the first-stage F statistics well above the standard rule of thumb of 10, below which concerns regarding weak instruments emerge.

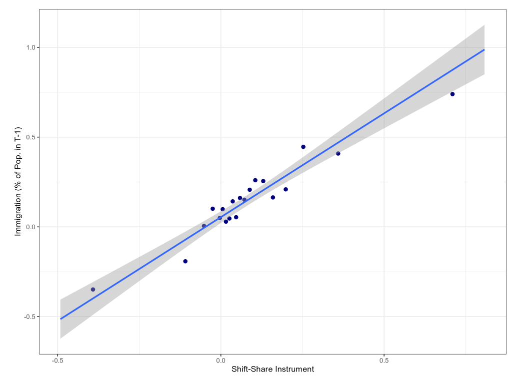
4.2 Headline Inflation

Table A3 presents the results from two-stage least squares estimation using the shift-share instrument using headline RPP inflation as the dependent variable. White (1980) heteroskedasticity-robust standard errors are reported in parentheses. The estimates indicate that overall immigration reduces local inflation relative to other MSAs (column 1). That this effect is almost doubled when looking only at the working-age population (col-

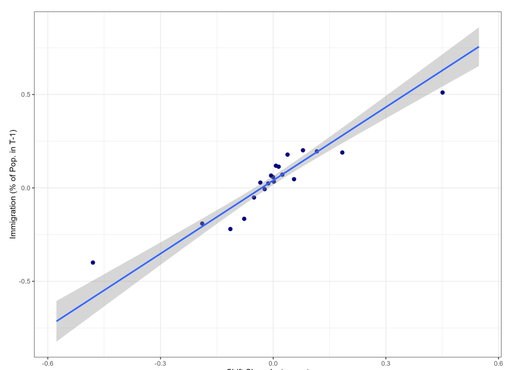
Figure 2: First Stage: Shift-Share IV



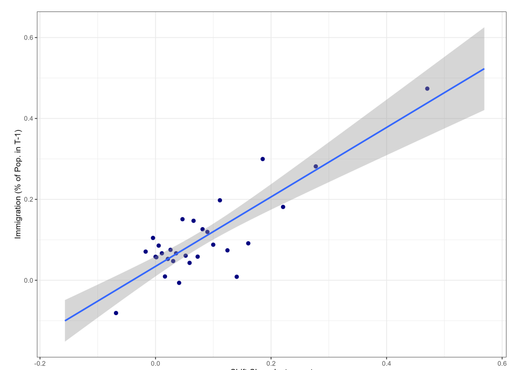
(a) All Residents



(b) Working Age: 18 to 64



(c) No College



(d) Some College

Notes: This figure presents binscatter plots of net immigration per capita on the shift-share instrument as defined in Equation 4. Panel a considers all immigrants, panel b considers all working age immigrants (Ages 18 to 64), panel c considers immigrants with no college education, and panel d considers immigrants with some college education. The data is from the American Community Survey.

umn 2), suggesting that labor supply may play an important role. Columns 3-6 attempt to tease out the effect of immigration by education level, and show that less educated immigrants in particular lower local prices. In contrast, more highly-educated immigrants have no discernable impact on inflation.

On average over our sample, annual migrant inflows are equivalent to around 0.2% of the population, a rate which approximately doubled in 2022. The marginal impact of such a doubling of immigration from the mean (or reducing immigration to zero) would lower (raise) inflation by about 0.1 to 0.2 p.p.[8] but with effects several times larger for immigrants with lower education levels. The impact of high-education immigration is insignificant.

The findings in Table A3 are also robust to controlling for US-born population growth, controlling for local unemployment, and restricting to the pre-COVID pandemic period (Table A5). We do not find evidence of spillovers to MSAs in the same state (Table A4).

4.3 Validity

4.3.1 Pre-trends

A threat to estimating the causal effect of net immigration on inflation is that trends of other variables, which might affect inflation, are correlated with changes in the instrument. We address this point by examining pre-trends as recommended by Goldsmith-Pinkham et al. (2020) and Borusyak et al. (2024).

We estimate our IV specification on inflation in lagged periods $T \in \{-1, -2, -3, -4\}$. Table 2 presents the results.[9] We find that our instrument is uncorrelated with inflation in the pre-period, with insignificant IV estimates on inflation in all lagged periods. We thus

[8] The impact of doubling or increasing immigration by 0.2 p.p. on inflation is $-0.657 \times 0.2 \simeq -0.13$. The total contribution of immigration at the doubled level (the 2022 level) or increasing immigration by 0.4 p.p. to inflation is $-0.657 \times 0.4 \simeq -0.3$.

[9] Pre-trends can also be seen visually in Figure A1.

Table 1: Main Results

	MSA-level Inflation (%)					
	All	18 to 64	No HS	No College	At least HS	Some College
	(1)	(2)	(3)	(4)	(5)	(6)
Immigration (% of Pop. in T-1)	-0.657^{***}	-1.096^{***}	-2.402^{**}	-2.032^{***}	-0.343	0.256
	(0.241)	(0.420)	(1.050)	(0.767)	(0.384)	(0.467)
IV F-statistic	42.66	26.62	27.81	21.19	27.62	37.04
Observations	3,588	3,588	3,588	3,588	3,588	3,588

*p<0.1; **p<0.05; ***p<0.01

Notes: This table presents estimates of Equation 1 using two-stage least-squares. The dependent variable is the log difference of the BEA's Implicit Regional Price Deflator which equals the product of the region's Regional Price Parity index and the U.S. PCE price index. Each observation is a MSA-Year pair. The shift-share instrument is defined in Equation 4 and uses data from the American Community Survey. Column 1 considers all immigrants, column 2 considers all working age immigrants (Ages 18 to 64), column 3 considers immigrants with no high school diploma, column 4 considers immigrants with no college education, column 5 considers immigrants with a high school diploma or higher,and column 6 considers immigrants with some college education. The sample covers 306 MSAs and years from 2009 to 2022. Robust standard errors are reported.

conclude that there are no differential pre-treatment trends across MSAs with different treatment (IV prediction) intensity.

4.3.2 Rotemberg weights

Following Goldsmith-Pinkham et al. (2020) we decompose the estimated effect β into the weighted average of estimates which use each origin country share ($sh_{c,i}$) as a *separate* instrument. That is:

$$\hat{\beta} = \sum_{c} \hat{\alpha}_c \hat{\beta}_c \tag{5}$$

where the weights $\hat{\alpha}_c$ are known as the "Rotemberg weights" after Rotemberg (1983), and the coefficients $\hat{\beta}_c$ comes from just-identified regressions using only origin country c's share as an instrument.

This decomposition permits two exercises which can help explain how variation in the data delivers the estimated coefficients. The first exercise is a simple report of the country-

Table 2: Pre-trends

	All	18 to 64	No HS	No College	At least HS	Some College
	(1)	(2)	(3)	(4)	(5)	(6)
A. Inflation (%) in $T = -1$						
Immigration (% of Pop. in T-1)	0.327	0.297	0.290	0.342	0.311	0.364
	(0.221)	(0.362)	(0.673)	(0.539)	(0.364)	(0.418)
B. Inflation (%) in $T = -2$						
Immigration (% of Pop. in T-1)	−0.111	−0.251	−1.133	−0.269	0.211	−0.433
	(0.250)	(0.411)	(0.952)	(0.607)	(0.429)	(0.489)
C. Inflation (%) in $T = -3$						
Immigration (% of Pop. in T-1)	−0.085	−0.131	−0.527	0.003	−0.071	0.196
	(0.252)	(0.362)	(0.714)	(0.534)	(0.423)	(0.497)
D. Inflation (%) in $T = -4$						
Immigration (% of Pop. in T-1)	−0.043	−0.040	0.782	0.147	−0.693	−0.504
	(0.225)	(0.304)	(0.579)	(0.343)	(0.518)	(0.431)
IV F-statistic	42.66	26.62	27.81	27.62	21.19	37.04
Observations	3,588	3,588	3,588	3,588	3,588	3,588

*p<0.1; **p<0.05; ***p<0.01

Notes: This table presents estimates of Equation 1 using two-stage least-squares. The dependent variable is the log difference of the Regional Price Parity index in years T =-1, -2, -3, and -4. Each observation is a MSA-Year pair. The shift-share instrument is defined in Equation 4 and uses data from the American Community Survey. Column 1 considers all immigrants, column 2 considers all working age immigrants (Ages 18 to 64), column 3 considers immigrants with no high school diploma, column 4 considers immigrants with no college education, column 5 considers immigrants with a high school diploma or higher,and column 6 considers immigrants with some college education. The sample covers 306 MSAs and years from 2009 to 2022. Robust standard errors are reported.

specific weights and estimates $\hat{\alpha}_c$ and $\hat{\beta}_c$. These are shown in Table 3 for the 5 origin country groups with the largest Rotemberg weights. No origin unduly dominates the estimates, with even the most influential contributor – Mexico, not surprisingly – attributed a Rotemberg weight of less than 25 percent. Moreover, although there is some variation in the point estimates, $\hat{\beta}_c$, they are consistently negative suggesting that our results are driven by broadly consistent effects across all origin regions, not just a handful of outliers.

We also use Table 3 as an opportunity to check the plausibility of concerns about reverse causality. To do this, we compute the share of outflows going to the most important MSA by country. If outflows from individual countries are highly concentrated among particular MSAs, MSA-specific variation in economic conditions might be plausibly driving outflows from specific countries. This would violate the exogeneity assumption of the instrument and invalidate a causal interpretation of our findings. However, emigrant flows appear to be well-distributed, with even the most important cities accounting for no more than one tenth of departing migrants for the three most influential countries.

Table 3: Rotemberg Weights

Origin	$\hat{\alpha}_c$	$\hat{\beta}_c$	Largest destination MSA	Share of emigration to largest MSA
Mexico	0.248	-0.774	Los Angeles-Long Beach-Anaheim, CA	0.08
Central America	0.184	-1.015	Houston-The Woodlands-Sugar Land, TX	0.09
India	0.131	-0.477	New York-Newark-Jersey City, NY-NJ-PA	0.10
West Indies	0.105	-0.822	New York-Newark-Jersey City, NY-NJ-PA	0.15
Cuba	0.085	-0.593	Miami-Fort Lauderdale-West Palm Beach, FL	0.22

Notes: This table reports the Rotemberg weights and point estimates for the weighted average decomposition of our estimate. Only the top five origin country regions are the top five origin countries according to the Rotemberg weights $\hat{\alpha}_c$. $\hat{\beta}_c$ is the coefficient from the just-identified regression. We report aggregated statistics, where we aggregate a given country across years as following the methodology in Goldsmith-Pinkham et al. (2020).

The second exercise that this calculation permits is an analysis of the correlation between the Rotemberg weights and the two components of our shift-share instrument: the shift and the share. We report the correlation matrix of the rotemberg weights α_c, aggregate national immigrant flows ΔF_c, and the variance in the initial shares sh_c across MSAs in

Table 4, again following the methodology in Goldsmith-Pinkham et al. (2020).[10] We find that our results are largely driven by the "shifts" – national immigrant flows – rather than the "shares". In particular, the shifts explain about 61% ($= 0.784^2$) of the variation in the weights, whereas the correlation between the weights and the variation in the origin country shares across MSAs is less strong.

Table 4: Rotemberg Weights: correlation with shifts and shares

	α_c	ΔF_c	Var(sh_c)
α_c	1		
ΔF_c	0.784	1	
Var(sh_c)	0.217	0.116	1

Notes: This table reports correlations between the weights, national immigrant inflows, and the variation in the origin country shares across MSAs. We report aggregated statistics, where we aggregate a given country across years as following the methodology in Goldsmith-Pinkham et al. (2020).

4.4 Heterogeneity by Goods vs Services

Beyond the impact on headline inflation, inward immigration may have a (potentially much larger) impact on relative prices. To investigate this issue, we repeat our analysis using the four major subcomponents of the RPP produced by the BEA. Table 5 reports the results of this exercise, which imply several main findings. First, that when immigration increases, goods inflation falls (with magnitudes similar to the overall price change) and utilities inflation increases. Moreover, for both categories, the effects are always largest for working age immigrants, are amplified when immigrants are less educated, and mitigated when immigrants are more educated. However, results for non-utilities, non-housing services are more mixed, showing no effect overall but disinflation in response to less educated immigrants and inflation when immigrants are more educated.

We also report results for housing inflation in Table 5, which we find an statistically null

[10] Specifically, we report aggregated statistics where we aggregate a given country across years: $\alpha_c = \sum_t \alpha_{c,t}$, $\beta_c = \sum_t \frac{\alpha_{c,t}}{\alpha_c} \beta_{k,t}$, and $\Delta F_c = \sum_t \frac{\alpha_{c,t}}{\alpha_c} |\Delta F_{c,t}|$.

response to inward immigration. However, the limitations of the RPP housing series discussed in Section 3 represent an important caveat on these findings. And so we present these results more for completeness than as convincing evidence on the impact on housing prices. Instead, we investigate housing costs more thoroughly in the next subsection and consider those our "headline" results for housing.

Table 5: Estimates by Inflation Category

	All	18 to 64	No HS	No College	At least HS	Some College
	(1)	(2)	(3)	(4)	(5)	(6)
A. Goods						
Immigration (% of Pop. in T-1)	-0.697^{**}	-1.064^{**}	-2.431^{**}	-1.689^{**}	-0.389	-0.327
	(0.284)	(0.461)	(1.068)	(0.746)	(0.468)	(0.503)
B. Services ex. Utilities and Housing						
Immigration (% of Pop. in T-1)	0.007	-0.117	-2.239^{**}	-1.125^{*}	0.792^{*}	1.311^{**}
	(0.246)	(0.378)	(1.136)	(0.638)	(0.445)	(0.565)
C. Services: Utilities						
Immigration (% of Pop. in T-1)	1.709^{**}	2.336^{*}	2.287	3.518^{**}	1.666	0.653
	(0.768)	(1.199)	(1.616)	(1.720)	(1.300)	(1.297)
D. Services: Housing						
Immigration (% of Pop. in T-1)	-1.119	-1.864	-0.496	-2.364	-1.721	-0.776
	(0.892)	(1.483)	(2.837)	(2.193)	(1.445)	(1.683)
IV F-statistic	42.66	26.62	27.81	21.19	27.62	37.04
Observations	3,588	3,588	3,588	3,588	3,588	3,588

*p< 0.1; **p< 0.05;***p< 0.01

Notes: This table presents estimates of Equation 1 using two-stage least-squares. The dependent variable is the log difference of the Regional Price Parity index by category as indicated in each panel. Each observation is a MSA-Year pair. The shift-share instrument is defined in Equation 4 and uses data from the American Community Survey. Column 1 considers all immigrants, column 2 considers all working age immigrants (Ages 18 to 64), column 3 considers immigrants with no high school diploma, column 4 considers immigrants with no college education, column 5 considers immigrants with a high school diploma or higher,and column 6 considers immigrants with some college education. The sample covers 306 MSAs and years from 2009 to 2022. Robust standard errors are reported.

4.5 Housing

Given our concerns about the RPP housing index, we consider two alternate measures of the cost of housing, the Zillow Home Value Index (ZHVI) and the Zillow Observed Rent

Index (ZORI). We discuss the relative merits of the Zillow indices versus the RPP housing series in detail in Section 3.[11]

Using these measures, we find that immigration increases local housing inflation relative to other MSAs. The effect is driven by high-education immigration (of at least high school graduates). On the other hand, low-education immigration (of non high school graduates) has no statistically significant impact on housing inflation. Table 6 presents the results.[12]

Table 6: Impact of Immigration on Housing Prices

	Housing Inflation (%): IV					
	All	18 to 64	No HS	No College	At least HS	Some College
Immigration (% of Pop.)	2.470***	3.319***	-0.991	2.842**	5.084***	4.284***
	(0.646)	(1.091)	(1.188)	(1.358)	(1.343)	(1.351)
IV F-statistic	42.66	26.62	27.81	27.62	21.19	37.04
Observations	3319	3319	3319	3319	3319	3319

$^{***}p < 0.01$; $^{**}p < 0.05$; $^{*}p < 0.10$

Notes: This table presents estimates of Equation 1 using two-stage least-squares. The dependent variable is the log difference of the Zillow Home Value Index (All Homes). Each observation is a MSA-Year pair. The shift-share instrument is defined in Equation 4 and uses data from the American Community Survey. Column 1 considers all immigrants, column 2 considers all working age immigrants (Ages 18 to 64), column 3 considers immigrants with no high school diploma, column 4 considers immigrants with no college education, column 5 considers immigrants with a high school diploma or higher,and column 6 considers immigrants with some college education. The sample covers 306 MSAs and years from 2009 to 2022. Robust standard errors are reported.

We find similar results when we look at housing rents instead of prices.[13] However, in contrast to purchase prices, rental inflation is driven more by those without a college degree. Table 7 presents the results.

[11]Reminder: the Zillow data 1) are corrected for changing composition at the individual unit level whereas the RPP adjustment is much cruder, 2) cover both rentals and owner-occupied housing versus just rentals in the RPP, and 3) considers current "market" prices whereas RPP measures prices with a lag due to long-term leases/rental contracts.

[12]Table A6 presents indices of other parts of the distribution for robustness.

[13]Rent data is only available from 2015 onwards, limiting our sample to 2015-2022.

Table 7: Impact of Immigration of Housing Rents

	Rental Inflation (%): IV					
	All	18 to 64	No HS	No College	At least HS	Some College
Immigration (% of Pop.)	2.076**	3.506*	1.604	3.316**	6.141	-0.138
	(1.040)	(1.940)	(1.187)	(1.566)	(4.595)	(1.415)
IV F-statistic	42.66	26.62	27.81	27.62	21.19	37.04
Num. obs.	1454	1454	1454	1454	1454	1454

$^{***}p < 0.01$; $^{**}p < 0.05$; $^{*}p < 0.10$

Notes: This table presents estimates of Equation 1 using two-stage least-squares. The dependent variable is the log difference of the Zillow Observed Rent Index. Each observation is a MSA-Year pair. The shift-share instrument is defined in Equation 4 and uses data from the American Community Survey. Column 1 considers all immigrants, column 2 considers all working age immigrants (Ages 18 to 64), column 3 considers immigrants with no high school diploma, column 4 considers immigrants with no college education, column 5 considers immigrants with a high school diploma or higher,and column 6 considers immigrants with some college education. The sample covers 306 MSAs and years from 2015 to 2022. Robust standard errors are reported.

4.6 Dynamic responses

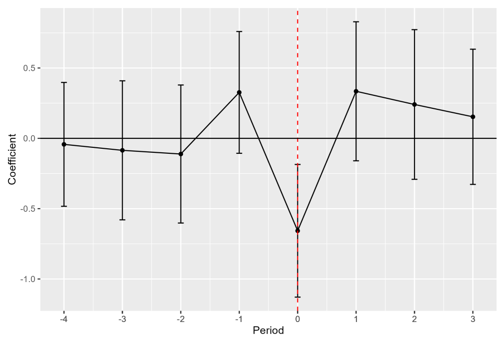
So far, our analysis has focused on the contemporaneous impact of immigration on local prices. This omits a possible important consequence of immigration, the dynamic response of prices over time. This is likely particularly important, especially for relative prices, if non-labor factors of production also adjust sluggishly to changes in immigrant labor supply.

First, we estimate the impact of immigration on the inflation rate. Following Jordà, 2005, we simply replace the dependent variable in equation (1) with the $t + h$ period equivalent separately for $h = 0, 1, 2, 3$. We present the results in Figure 3a.[14] We find that immigration only lowers contemporaneous inflation but it raises it thereafter, although the effects are statistically insignificant.

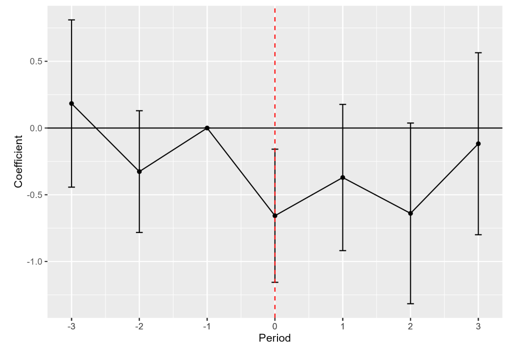
Next, we analyze the dynamic response of the price level (rather than inflation rate) by calculating the local projection equivalents of our contemporaneous estimates for equation (1). In this case, the dependent variable is the change in the price level in period $t + h$ relative to period $t - 1$. We present the results in Figure 3b. Again, we find that

[14]We present the results by immigrant demographic group in Figure A1.

Figure 3: Dynamic responses



(a) Inflation Rate



(b) Price level

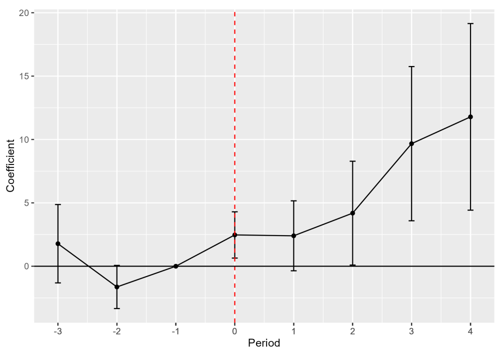
Notes: This figure presents estimates of Equation 1 using two-stage least-squares for lags/leads of the dependent variable from periods -4 *to* 3*. In panel a, the dependent variable is the log difference of the BEA's Implicit Regional Price Deflator between period t and period t-1. In panel b, the dependent variable is the log difference of the BEA's Implicit Regional Price Deflator between period t and period -1. Each observation is a MSA-Year pair. The shift-share instrument is defined in Equation 4 and uses data from the American Community Survey.*

immigration only lowers the overall price level contemporaneously, and the price level is statistically indistinguishable from that in $t-1$ in future years.

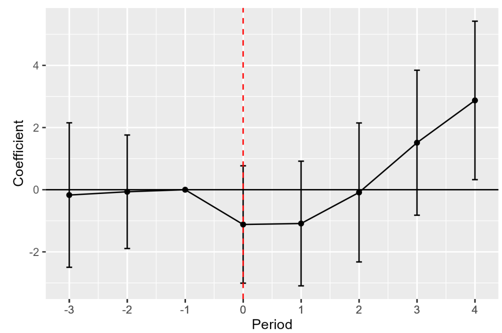
Except for housing, the price level impact of immigration on all sub categories of inflation is transitory.[15] We find that immigration increases housing prices persistently over time (positive and statistically significant impact as far as 4 years), driven by high-education immigration. And although, low-education immigration has no impact on housing prices in the first year, low-education immigration lowers housing prices persistently starting in the second year (negative and statistically significant impact as far as 3 years). This is consistent with increased (low-education) labor supply to housing construction impacting housing prices with a lag, a point we discuss further in Section 5. Figure 4 presents the results for both the RPP housing index and the Zillow index.

[15]See Appendix Figure A2.

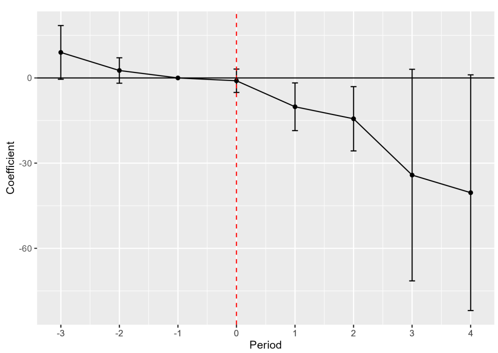
Figure 4: Impact of Immigration of Housing Prices (% change from $T = -1$)



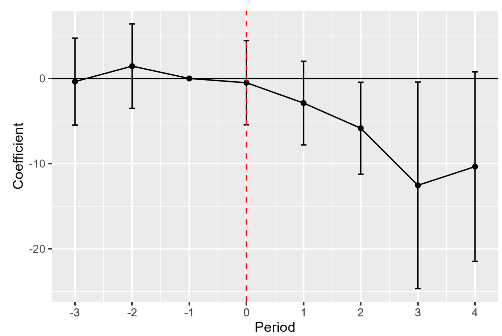
(a) Zillow: All Residents



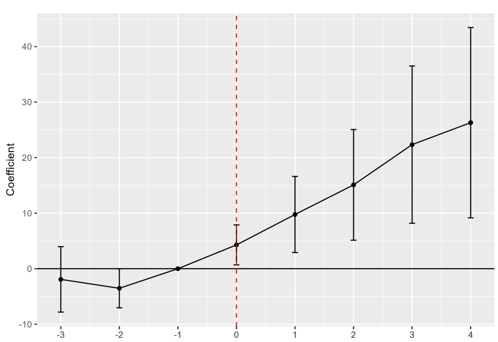
(b) RPP: All Residents



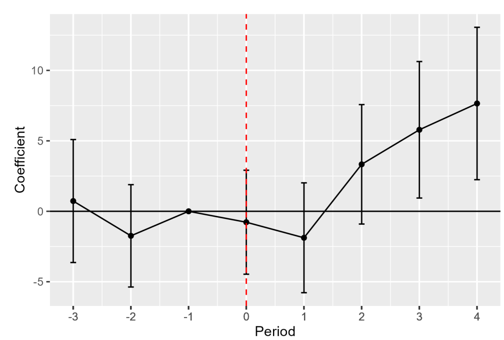
(c) Zillow: No HS Diploma



(d) RPP: No HS Diploma



(e) Zillow: Some College



(f) RPP: Some College

Notes: This figure presents estimates of Equation 1 using two-stage least-squares for lags/leads of the dependent variable from periods -3 *to* 4*. The dependent variable is the log difference of the Zillow Home Value Index (All Homes) or the Regional Price Parities index for housing. Each observation is a MSA-Year pair. The shift-share instrument is defined in Equation 4 and uses data from the American Community Survey. Panel a considers all immigrants, panel b considers immigrants with no high school diploma, panel c considers immigrants with no college education, and panel d considers immigrants with some college education.*

4.7 Comparison to other estimates

Comparing our estimated effects to those in the literature suggests that the horizon over which immigration has an impact on prices is very important. For example, Cortes (2008) uses decennial census data to compute an impact of increases in low-skilled immigrants on prices which, when converted to equivalent units to ours, would be equivalent to a coefficient of roughly -0.09.[16] This is about twenty times smaller than our most directly comparable estimates, those for low-education workers (between -2.0 and -2.4), and suggests that the long run estimates miss important contemporaneous effect which might fade. Indeed, our results on the dynamic response of prices suggest that this is in fact true.

In contrast, Lach (2007) estimates the contemporaneous response of prices to immigration from the former Soviet Union to Israel in 1990 using purely cross-sectional variation. He finds that a one percentage point increase in the immigrant-to-native ratio reduced grocery store goods prices by 0.5 percent point, a number very close to our findings for goods (see Table 5). It is also worth emphasizing that our effective sample size is much greater than either of Cortes (2008) or Lach (2007). Although both papers enrich their sample by pooling prices across different product categories, variation in immigration is only at city or city-decade level, giving only 90 and 52 distinct values respectively for the dependent variable. In contrast, we have over 3500 distinct values in our study, giving us increased statistical power.

[16] Cortes (2008) finds that a 10 percent increase in the total number of low-skilled immigrants over a decade reduces average prices of immigrant-intensive services by 2 percent. Since these have an expenditure share of around 3.5 percent, and effects on other components are negligible, this implies a $-2 \times 0.035 \simeq -0.07$ impact on the price level over ten years, or approximately 0.007 per year. A similar simple calculation underlies her counterfactual simulations. Since the foreign-born share of the US population during 1980-2000 is around 8 percent, a ten percent increase in the immigrant population over a decade is equivalent to a 0.08pp yearly increase in migrant inflows relative to population. Thus, the annual per-percentage impact is $-0.007 \div 0.08 \simeq -0.09$. Albeit crude, this calculation is at least a useful approximation for orders of magnitude.

5 Interpretation

In the preceding section we outlined three key results. One is about the overall price level: when a local area experiences more immigration, inflation is lower relative to elsewhere. Another is about the response of product-specific local prices to the same shock: goods prices fall, services change little, and utility and housing increase (although the latter depends somewhat on the measure used). The last result is about the type of immigrants. We find that the magnitude of the response of local prices to immigration is larger when immigrants have less eduction.

In this section we present a simple framework to make sense of these results collectively. The key insight is that the direction of (relative) local price changes is a function of the relative elasticities of supply of different product, and that the magnitude of those changes depends on consumers' elasticity of substitution between them. To focus on the intuition, our exposition is intentionally heuristic, and we offer only graphical support for our arguments. In the Appendix we present a mathematical representation of the same in which we can derive a similar result algebraically in closed form.

We consider a region producing two goods, x and y.[17] Consumers have convex and homothetic preferences over both types of good. For simplicity, we consider the simplest case where all consumption is produced locally. The key ingredient which captures the idea that relative supply and demand can be affected differently by increased immigration is non-homothetic production. Specifically, we allow the two goods to have different elasticities of supply. Such a difference in elasticities of supply might arise for any number of reasons. This difference in supply elasticities gives rise to a change in relative prices when labor supply expands, even if preferences are homothetic (and so the increase in demand due to immigration is not biased towards any one product).

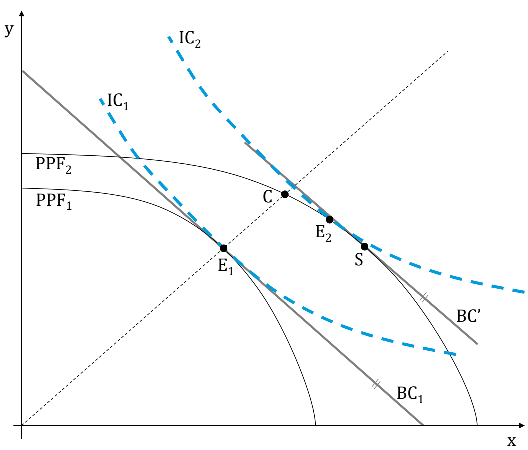
[17] Here we use "goods" as a stand-in for a category of consumption, as is common. this is not the sense in which we use "Goods" as opposed to services in our empirical work.

To see this more clearly, Figure 5 presents a graphical analysis of this framework. The region has a production possibility frontier given initially by PPF_1. Consumers' preferences define an efficient allocation E_1 with corresponding indifference curve IC_1. In a decentralized equilibrium, prices move such that markets clear and the budget constraint BC_1 is tangent to both PPF_1 and IC_1 at E_1.[18] An increase in immigration increases the availability of labor, expanding the set of production possibilities. But this expansion is uneven, with one good increasing production relatively more for a given increase in inputs. That is, one of the goods has a more elastic supply. This could be a function of the intrinsic production technologies for the two sectors, for example because of differing ease of substituting labor for fixed factors such as capital. Or it could be because of specialization of immigrant labor for one of the goods. However, even in this case, the PPF also shifts out for both goods so long as some domestic workers can reallocate. Figure 5a illustrates the shift in the production possibility frontier case where product x is more elastic, and Figure 5b when product y is more elastic.

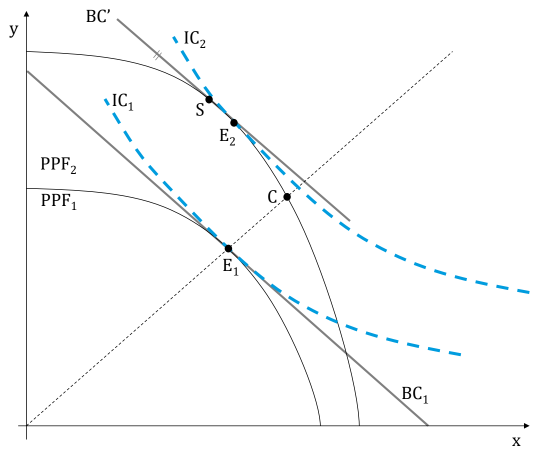
Even with homothetic preferences, production shifts towards the more elastically supplied good (point E_2 in Figure 5). The extent of this shift, though, depends on the willingness of households to switch between the two products, i.e. their elasticity of substitution. To see this, consider first the extreme of perfect complements. Then the efficient allocation features production in proportion to the initial allocation (point C). But if the two goods were perfect substitutes in consumption, the new efficient allocation would be the point on PPF_2 with tangent parallel to the tangent to PPF_1 at the original allocation. This is given by S. Because the production possibility frontier expands more towards the more

[18] With convex level sets for preferences and production, the second welfare theorem holds and any efficient allocation can be implemented as a decentralized equilibrium where firms and households are price-takers. However, finite elasticity of supply implies a non-constant returns to scale technology, precluding the use of a representative firm. To avoid having to deal with a distribution of firm sizes in this simple example, we focus on the efficient allocation and appeal to the second welfare theorem to argue that a decentralized equilibrium implementing this could be found, even though we do not go through the details here. In the Appendix, we make this connection explicit, showing how to rationalize an upward-sloping short-run supply curve for each product even when firms are price-takers, using a production technology with product-specific substitution of labor and capital and fixed capital in the short run.

Figure 5: Local impact of increased immigration with non-homothetic production: a simple model



(a) x is relatively elastic



(b) y is relatively elastic

In both figures PPF_1 and PPF_2 are the production possibility frontiers, before and after immigration respectively. The homothetic indifference curves IC_1 and IC_2 then define efficient allocations E_1 and E_2. BC_1 is the budget constraint which would support E_1 in a decentralized equilibrium. However, BC' is *not* the budget constraint which would support E_2. Instead, it is the tangent of PPF_2 parallel to BC_1, and so meets PPF_2 at S, the efficient allocation if tradeables and nontradeables were perfect substitutes. The point labelled C is the other extreme, and would be the efficient allocation if the two consumption goods were perfect complements.

elastically supplied good, this tangency condition is satisfied only by further increasing production of the more elastically supplied product. The change in relative prices moves opposite to the change in quantities.

In the Appendix, we write down the equilibrium conditions for this model mathematically and compute the equations for decentralized equilbrium and derive a closed-from solution for the derivative of relative prices in the decentralized equilibrium with respect to labor supply in one of the regions. This can be decomposed into two parts:

$$\frac{d\log(p_y/p_x)}{d\log L_1} = \underbrace{\left(\frac{\epsilon_x}{1+\epsilon_x} - \frac{\epsilon_y}{1+\epsilon_y}\right)}_{\text{Direct effect}} \underbrace{(1+\gamma)}_{\text{Indirect effect}}$$

where ϵ_x, ϵ_y are the elasticities of supply of goods x and y respectively, and γ a deceasing function of the elasticity of substitution, which is positive if and only if the elasticity of substitution is smaller than one, and is -1 in the case of perfect substitutes. The direct effect captures the change in price if labor shares in the two industries remained constant. The indirect effect arises because this change in relative prices induces a reallocation of labor, towards the cheaper good if they are substitutes and away from it if they are complements. In this sense, the elasticity of substitution stands in for differences in the elasticity of demand for the two products. The response of the aggregate price level is then just the average of the product-specific price levels weighted by their relative expenditure shares.

While simple, we think that this setting offers some useful insight into how immigration might affect local price levels in equilibrium. The central point is that the relative elasticity of supply determines changes in the local relative price of different goods. Goods that are relatively more elastically supplied should experience relatively less inflation. But is this consistent with our empirical results? We argue that it is. That labor and capital are closer substitutes in production of goods rather than services is long-established.[19] If capital

[19] Including, among others Hamermesh, 1996, Herrendorf et al., 2015, Alvarez-Cuadrado et al., 2018, and Vom Lehn, 2020. Intuitively, an industry being more capital-intensive tends to be a symptom of capital and

is fixed in the short term and labor becomes more readily available, then the supply of goods will be more elastic than that of services, since services producers cannot easily increase capital to complement increased labor usage. This is consistent with our finding that the relative price of goods drops by more than by services. Likewise, it seems highly plausible that the supply of housing is very inelastic in the short term – for example, if stocks of land or construction materials are in fixed supply in the short term – driving up prices.

A similar logic applies to our results on education level. A large literature[20] establishes that less-educated labor is typically more substitutable for capital. Again, this is would predict a more elastic supply of the types of goods produced by less-educated workers, and pushing down their relative prices. That the differences between the responses of the broad categories of goods are relatively large is consistent with them being complementary, as one would expect.

Finally, this simple framework can also help make sense of our dynamic results. The assumption of fixed capital is plausible only in the short run. If the supply of capital is infinitely elastic in the long-run, which seems like a reasonable assumption at the MSA level, then the long-run expansion in the production possibility frontier will be homothetic, leaving relative prices unaffected. This is consistent with our findings that the long run impact on relative price levels is indistinguishable from zero for goods, services, and utilities. The only exception is housing, and industry where land is in fixed supply even in the long run.

Of course, the real world departs from our simple model in many regards. For example, relative price levels across regions will be linked by the trade across them. And if immigrants are numerous enough with sufficiently different preferences to native workers,

labor being more substitutable. This is because labor is the relatively scare factor, so wages increase over time relative to capital costs. Accordingly, firms replace labor with capital over the long run, just more so in industries where inputs are closer substitutes.

[20]See Griliches, 1969, Krusell et al., 2000, Chen, 2020, and many others.

then we cannot reasonably model aggregate preferences as homothetic. But the consistency of our results with an even simpler framework shows that we can at least give a coherent interpretation to our results without resort to these other channels.

6 Conclusion

In this paper, we study how immigration impacts local inflation. Using a shift-share instrument, we find that immigration reduces local inflation relative to other MSAs, with effects most pronounced when immigrants are working-age and have less educations. We also find that immigration lowers local goods inflation, increases housing inflation, and has a negligible effect on services. However, for all products, less educated immigrants tend to be more disinflationary than more educated immigrants. Longer-run price effects are much smaller and typically statistically indistinguishable from zero. One exception is the price of housing, where the (dis)inflationary effects of (low-) high-education immigrants can last several years.

We offer a simple interpretation for our results, arguing that they are consistent with inward immigration constituting an increase in labor supply when production is non-homothetic in the short turn. We motivate such non-homotheticities with the observation that other factors, most notably capital, are fixed in the short run. This simple model makes an equally simple prediction: relative price declines should be largest in industries (and for immigrant labor types types) where labor is more substitutable for capital. This is exactly what we see in our results. Goods prices decline, housing increases; less-educated immigrants lower prices, and higher-educated ones raise them. This simple framework also makes a prediction about the longer run, when capital is not fixed and so non-homotheticities are less important. This is consistent with our finding that long-run prices responses are statistically indistinguishable from zero. That the one exception to this finding is an industry with another fixed factor – housing, where land is an essential input – reinforces our interpretation. There are of course, other interpretations for our findings, but likely none simpler, something we see as a virtue.

Some other related issues remain unaddressed. For example, more detailed pricing data could shine light on the specific goods and services most affected by immigration. More recent data on surveys of consumer views could investigate the extent to which opinions on local prices react to immigration, and how they differ from reality. And the simple mechanisms we illustrate in our static framework could be extended to a full dynamic stochastic general equilibrium model. We leave these issues for further research.

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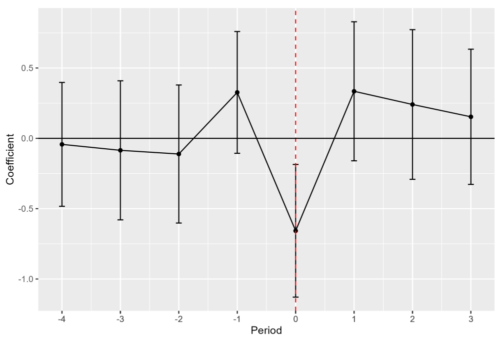
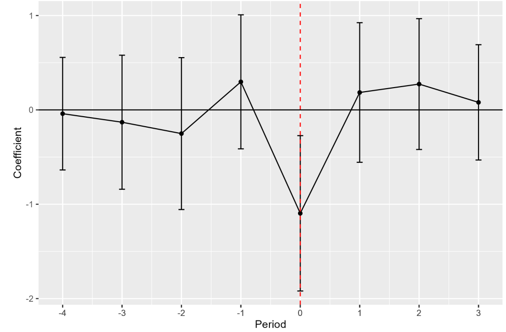
A Appendix Figures and Tables

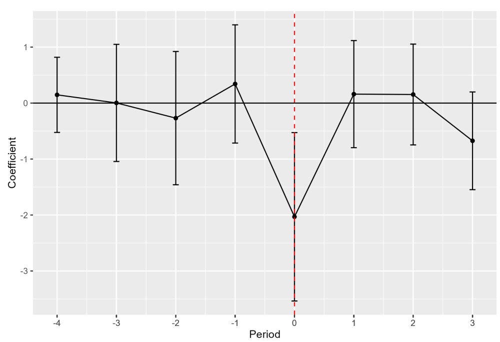
Figure A1: Dynamic responses – Inflation



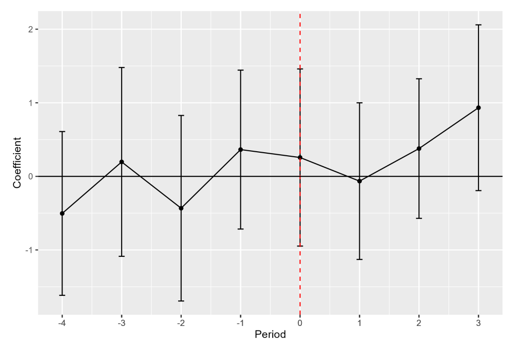
(a) All Residents



(b) Working Age: 18 to 64



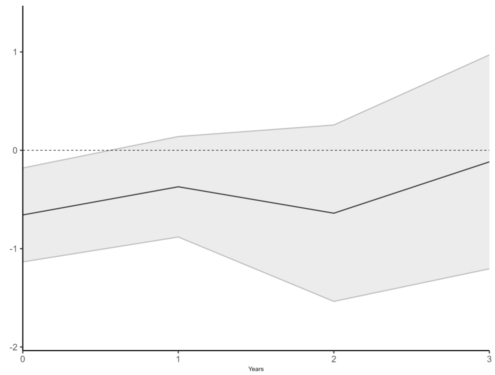
(c) No College



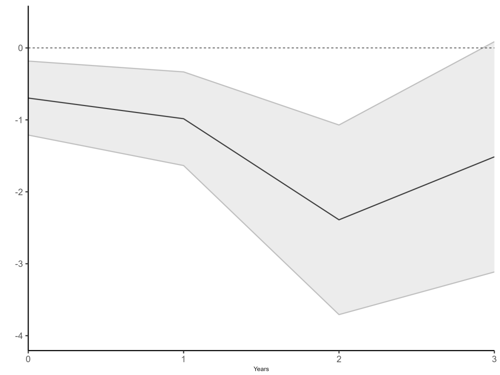
(d) Some College

Notes: This figure presents estimates of Equation 1 using two-stage least-squares for lags of the dependent variable from periods -4 *to* 4*. The dependent variable is the log difference of the BEA's Implicit Regional Price Deflator which equals the product of the region's Regional Price Parity index and the U.S. PCE price index. Each observation is a MSA-Year pair. The shift-share instrument is defined in Equation 4 and uses data from the American Community Survey. Panel a considers all immigrants, panel b considers all working age immigrants (Ages 18 to 64), panel c considers immigrants with no college education, and panel d considers immigrants with some college education.*

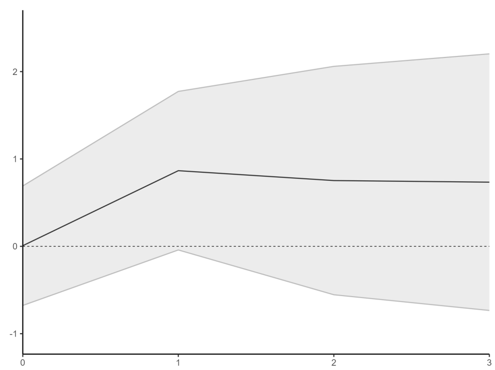
Figure A2: Impulse Responses



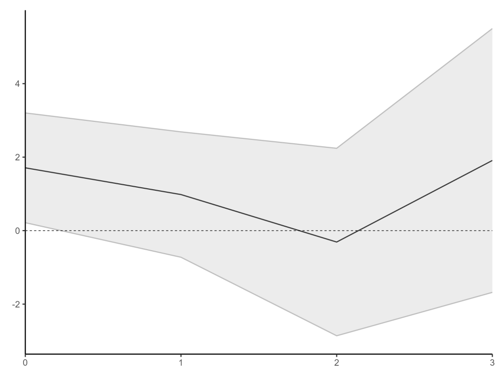
(a) Overall Inflation



(b) Goods



(c) Services ex. Housing and Utilities



(d) Utilities

This figure presents impulse responses following Jordà (2005) and Adämmer (2019) for a 1 p.p. shock to net immigration per capita. The dependent variable is log Regional Price Parities by category, and the explanatory variable is net immigration divided by the population in T-1. Panel (a) reports results for headline inflation, panel (b) reports results for goods, panel (c) reports results for other services, and panel (d) reports results for utilities. Each observation is a MSA-Year pair. The shift-share instrument is defined in Equation 4 and uses data from the American Community Survey.

Table A1: Summary Statistics

Variable	N	Mean	Std. Dev.	Pctl. 25	Median	Pctl. 75
Inflation (%)	3614	1.8	2.6	0.17	1.5	3.2
Population	3614	991411	1938853	175219	377469	860540
Foreign-born Share	3614	0.11	0.077	0.055	0.084	0.14
Age 18 to 64 Share	3614	0.62	0.034	0.6	0.62	0.63
At least High School Share	3614	0.56	0.048	0.54	0.56	0.58
Some College Share	3614	0.33	0.062	0.29	0.33	0.37
Net Immigration per cap.	3614	0.0019	0.012	-0.0044	0.0018	0.0082
Net Immigration (Age 18-64) per cap.	3614	0.0012	0.01	-0.0039	0.0014	0.0065
Net Immigration (At least HS) per cap.	3614	0.0015	0.0089	-0.0032	0.0015	0.0061
Net Immigration (Some college) per cap.	3614	0.00095	0.0067	-0.0026	0.00096	0.0044

Notes: This table presents summary statistics of the data. Inflation data comes from the BEA's Regional Price Parity index and the population data comes from the American Community Survey. Each observation is a MSA-Year pair.

Table A2: First Stage: Shift-Share IV

	Immigration (% of Pop. in T-1)					
	All	18 to 64	No HS	No College	At least HS	Some College
	(1)	(2)	(3)	(4)	(5)	(6)
Shift-Share Instrument	0.616***	0.568***	0.449***	0.568***	0.536***	0.796***
	(0.087)	(0.096)	(0.164)	(0.123)	(0.096)	(0.103)
Observations	4,443	4,443	4,443	4,443	4,443	4,443
IV F-statistic	42.66	26.62	27.81	21.19	27.62	37.04

*p<0.1; **p<0.05; ***p<0.01

Notes: This table presents regression estimates of net immigration per capita on the shift-share instrument as defined in Equation 4. Column 1 considers all immigrants, column 2 considers all working age immigrants (Ages 18 to 64), column 3 considers immigrants with no high school diploma, column 4 considers immigrants with with no college education, column 5 considers immigrants with a high school diploma or higher, and column 6 considers immigrants with some college education. The data is from the American Community Survey. Robust standard errors are reported.

Table A3: Impact of Immigration on Inflation – Alternative Education Groups

	MSA-level Inflation (%)				
	All	18 to 64	No HS	Only HS	Some College
	(1)	(2)	(3)	(4)	(5)
Immigration (% of Pop. in T-1)	-0.657^{***}	-1.096^{***}	-2.402^{**}	-1.728^{*}	0.256
	(0.241)	(0.420)	(1.050)	(0.971)	(0.467)
IV F-statistic	42.66	26.62	27.81	10.94	37.04
Observations	3,588	3,588	3,588	3,588	3,588

*p<0.1; **p<0.05; ***p<0.01

Notes: This table presents estimates of Equation 1 using two-stage least-squares. The dependent variable is the log difference of the BEA's Implicit Regional Price Deflator which equals the product of the region's Regional Price Parity index and the U.S. PCE price index. Each observation is a MSA-Year pair. The shift-share instrument is defined in Equation 4 and uses data from the American Community Survey. Column 1 considers all immigrants, column 2 considers all working age immigrants (Ages 18 to 64), column 3 considers immigrants with no high school diploma, column 4 considers immigrants with only a high school diploma, and column 6 considers immigrants with some college education. The sample covers 306 MSAs and years from 2009 to 2022. Robust standard errors are reported.

Table A4: Impact of Immigration on Inflation – Spillovers

	MSA-level Inflation (%)					
	All	18 to 64	No HS	No College	At least HS	Some College
	(1)	(2)	(3)	(4)	(5)	(6)
MSA Immigration (% of Pop. in T-1)	-0.721^{**}	-1.158^{**}	-3.557	-2.004^{**}	-0.480	-0.060
	(0.320)	(0.526)	(3.025)	(0.891)	(0.493)	(0.572)
State Immigration (% of Pop. in T-1)	0.156	0.153	1.632	-0.049	0.388	1.015
	(0.422)	(0.713)	(2.779)	(0.799)	(0.732)	(0.939)
IV F-statistic	21.96	13.33	16.66	14.14	10.6	18.66
Observations	3,588	3,588	3,588	3,588	3,588	3,588

*p<0.1; **p<0.05; ***p<0.01

Notes: This table presents estimates of Equation 1 using two-stage least-squares with an additional explanatory variable – total immigration as a percentage of population in $T-1$ in the state or states which the MSA is in – and an additional instrument – predicted total immigration as a percentage of population in $T-1$ in the state or states which the MSA is in. The dependent variable is the log difference of the BEA's Implicit Regional Price Deflator which equals the product of the region's Regional Price Parity index and the U.S. PCE price index. Each observation is a MSA-Year pair. The shift-share instrument is defined in Equation 4 and uses data from the American Community Survey. Column 1 considers all immigrants, column 2 considers all working age immigrants (Ages 18 to 64), column 3 considers immigrants with no high school diploma, column 4 considers immigrants with no college education, column 5 considers immigrants with a high school diploma or higher, and column 6 considers immigrants with some college education. The sample covers 306 MSAs and years from 2009 to 2022. Robust standard errors are reported.

Table A5: Impact of Immigration on Inflation – Robustness

	All	18 to 64	No HS	No College	At least HS	Some College
	(1)	(2)	(3)	(4)	(5)	(6)
A. Restrict to pre-pandemic period (before 2020)						
Immigration (% of Pop. in T-1)	-0.476^{*}	-0.750^{**}	-3.314^{*}	-1.470^{**}	-0.248	-0.216
	(0.243)	(0.351)	(1.796)	(0.647)	(0.351)	(0.496)
Observations	2,828	2,828	2,828	2,828	2,828	2,828
B. Control for % growth of the US-born population						
Immigration (% of Pop. in T-1)	-0.724^{***}	-1.312^{**}	-2.432^{**}	-2.132^{***}	-0.341	0.256
	(0.264)	(0.531)	(1.060)	(0.818)	(0.381)	(0.465)
C. Control for MSA Unemployment Rate						
Immigration (% of Pop. in T-1)	-0.759^{***}	-1.309^{***}	-3.362^{**}	-2.681^{***}	-0.405	0.283
	(0.238)	(0.435)	(1.517)	(1.000)	(0.361)	(0.440)
D. Control for Lagged MSA Unemployment Rate						
Immigration (% of Pop. in T-1)	-0.760^{***}	-1.324^{***}	-3.552^{**}	-2.741^{***}	-0.369	0.298
	(0.238)	(0.439)	(1.636)	(1.028)	(0.350)	(0.434)
Observations	3,588	3,588	3,588	3,588	3,588	3,588

$^{*}p<0.1$; $^{**}p<0.05$; $^{***}p<0.01$

Notes: This table presents estimates of Equation 1 using two-stage least-squares. The dependent variable is the log difference of the Regional Price Parity index. Each observation is a MSA-Year pair. The shift-share instrument is defined in Equation 4 and uses data from the American Community Survey. Column 1 considers all immigrants, column 2 considers all working age immigrants (Ages 18 to 64), column 3 considers immigrants with no high school diploma, column 4 considers immigrants with no college education, column 5 considers immigrants with a high school diploma or higher,and column 6 considers immigrants with some college education. The sample covers 306 MSAs, and years from 2009 to 2019 for Panel A and 2009 to 2022 for Panel B. Panel A includes a control for percentage growth of the US-born population. Panel C includes a control for the MSA unemployment rate in year T. Panel D includes a control for the MSA unemployment rate in year T-1. Robust standard errors are reported.

Table A6: Impact of Immigration of Housing Prices - By Zillow Series

	Housing Inflation (%): IV					
	All	18 to 64	No HS	No College	At least HS	Some College
A. Top-tier Zillow Home Value Index: 65th to 95th percentile range						
Immigration (% of Pop.)	2.600***	3.684***	0.780	3.743***	4.857***	4.152***
	(0.635)	(1.094)	(1.044)	(1.427)	(1.283)	(1.279)
B. Zillow Home Value Index: 35th to 65th percentile range						
Immigration (% of Pop.)	2.470***	3.319***	-0.991	2.842**	5.084***	4.284***
	(0.646)	(1.091)	(1.188)	(1.358)	(1.343)	(1.351)
C. Bottom-tier Zillow Home Value Index: 5th to 35th percentile range						
Immigration (% of Pop.)	3.912***	5.106***	-1.740	4.153**	7.970***	6.408***
	(0.874)	(1.490)	(1.528)	(1.768)	(1.926)	(1.679)
IV F-statistic	42.66	26.62	27.81	27.62	21.19	37.04
Observations	3319	3319	3319	3319	3319	3319

***$p < 0.01$; **$p < 0.05$; *$p < 0.10$

Notes: This table presents estimates of Equation 1 using two-stage least-squares. The dependent variable is the log difference of the Zillow Home Value Index (All Homes) by standard, top-tier and bottom-tier. Each observation is a MSA-Year pair. The shift-share instrument is defined in Equation 4 and uses data from the American Community Survey. Column 1 considers all immigrants, column 2 considers all working age immigrants (Ages 18 to 64), column 3 considers immigrants with no high school diploma, column 4 considers immigrants with no college education, column 5 considers immigrants with a high school diploma or higher,and column 6 considers immigrants with some college education. The sample covers 306 MSAs and years from 2009 to 2022. Robust standard errors are reported.

B Simple model

B.1 Set up

B.1.1 Technology

A region produces two goods, x and y, which we think of as wholly consumed locally.[21] The aggregate production technology for both goods is decreasing returns to scale using only labor as an input, with:

$$X = AL_x^{\alpha} \tag{6}$$

$$Y = BL_y^{\beta} \tag{7}$$

The different curvatures $\alpha, \beta < 1$ of the production function introduce different marginal cost curves and hence different elasticities of supply. In Section B.3 we justify this aggregate production technology as an approximation to a situation with a large number of price-taking firms each using a CES technology combining labor and capital but where capital is fixed in the short term. But for now we take this as given.

We assume perfect competition[22] so price equals marginal cost:

$$P_x = \frac{1}{\alpha} W A^{-1/\alpha} X^{1/\alpha - 1} \tag{8}$$

$$P_Y = \frac{1}{\beta} W B^{-1/\beta} Y^{1/\beta - 1} \tag{9}$$

Where W is the common wage across sectors.

Then the firm makes profits, with markups $\frac{1}{\alpha} - 1$ and $\frac{1}{\beta} - 1$ in the two sectors, which we assume is rebated to households.

[21] A natural extension is to the case where there is trade between regions and x and y are tradeables and nontradeables. However, data on local tradeables prices are not readily available.

[22] Again, we justify this with our more detailed microfoundation in Section B.3.

B.1.2 Preferences

There are L households in the region, and each household inelastically supplies a unit of labor. Preferences are CES across the two goods. Thus households solve:

$$\max_{X,Y} \; \mathcal{U}_i = \left(\eta_x X^{\rho} + \eta_y Y^{\rho}\right)$$
$$\text{s.t. } \; P_X X + P_Y Y = (1+\theta) WL$$

where and total household income is labor income WL plus firms' profits, where θ is the average markup across the two sectors.

$$\Theta = \Lambda \left(\frac{1}{\alpha} - 1\right) + (1-\Lambda)\left(\frac{1}{\beta} - 1\right) \tag{10}$$

Letting $\sigma = 1/(1-\rho)$, the households' demand curves are:

$$X = \left(\frac{\eta_x}{P_x}\right)^{\sigma} \frac{\theta WL}{P^{1-\sigma}} \tag{11}$$
$$Y = \left(\frac{\eta_y}{P_y}\right)^{\sigma} \frac{\theta WL}{P^{1-\sigma}} \tag{12}$$

Where:

$$P^{1-\sigma} = \eta_x^{\sigma}(P_X)^{1-\sigma} + \eta_y^{\sigma}(P_Y)^{1-\sigma} \tag{13}$$

B.1.3 Equilibrium

We close the model with zero excess demand in the labor market and balanced trade.

$$L_X = \Lambda L \tag{14}$$

$$L_Y = (1 - \Lambda)L \tag{15}$$

Where Λ is the fraction of labor in industry Y.

Taking A, B, L and all the other parameters as given, equilibrium is then values:

$$\{X, Y, L_X, L_Y, P, P_X, P_Y, W, \Lambda, \Theta\}$$

satisfying equations (6)-(15). This is 10 equations in 10 unknowns.[23]

B.2 Differential form

We now differentiate the above equations w.r.t. L in order to get the marginal change in equilibrium with respect to increased population in the smaller region.

Notation: lower case variables will denote the log-log derivative, for example:

$$y_i = \frac{d \log Y_i}{d \log L_1}$$

[23]That said, one equation is redundant in that the overall price level is not determined without a numeraire assumption. For any given solution, a doubling of all prices and wages will also be a solution.

Eliminating L_X and L_Y gives us a system of linear simultaneous equations:

$$x = \alpha\lambda + \alpha \tag{16}$$

$$y = -\beta\Lambda^*\lambda + \beta \tag{17}$$

$$p_x = w + \left(\frac{1-\alpha}{\alpha}\right)x \tag{18}$$

$$p_y = w + \left(\frac{1-\beta}{\beta}\right)y \tag{19}$$

$$y = -\sigma p_y + \theta + w - (1-\sigma)p + 1 \tag{20}$$

$$x = -\sigma p_x + \theta + w - (1-\sigma)p + 1 \tag{21}$$

$$\theta = \Theta^*\lambda \tag{22}$$

$$p = s_x p_x + s_y p_y \tag{23}$$

Where $\Lambda^* = \Lambda/(1-\Lambda)$, $\Theta^* = (\beta-\alpha)/(\beta(1-\alpha)\Lambda + \alpha(1-\beta)(1-\Lambda))$ and s_x and s_y are the expenditure shares of X and Y respectively.

Then, we can solve for λ:

$$\lambda = \frac{-(\alpha-\beta)(1-\sigma)}{\alpha + \beta\Lambda^* + \sigma\mu}$$

Where $\mu = (1-\alpha) + \Lambda^*(1-\beta)$.

Then solving for the relative price level gives us:

$$p_y - p_x = (\alpha-\beta)\left(1 + \frac{(1-\sigma)\mu}{\alpha + \beta\Lambda^* + \sigma\mu}\right)$$

Since $\epsilon_x = \alpha/(1-\alpha)$ and $\epsilon_y = \beta/(1-\beta)$ are the elasticities of supply for X and Y respectively, this satisfies the form in the main paper with $\gamma = (1-\sigma\mu)/(\alpha + \beta\Lambda^* + \sigma\mu)$.

B.3 Elasticity of supply and substitability of labor for capital

One shortcoming of the simple example above is that it just imposes a decreasing returns to scale (DRS) aggregate production technology. Since DRS technologies do not, in general, aggregate cleanly like constant returns to scale ones do, we cannot justify this with a representative firm using a production technology identical to the aggregate. Instead, we show that – to a first-order approximation – this can be rationalized by a large number of small firms using a constant elasticity of substitution production function using capital and labor as inputs but where capital is fixed at the firm level.

B.3.1 Flexible factor usage

For reference, we start with the case where both inputs are variable. This is the textbook case, but is a useful model for what follows when we fix one input in the following section. Consider an individual firm i which produces output from labor and capital taking factor prices as given. Then they face the cost minimization problem:

$$\min_{l_i,k_i} wl_i + rk_i$$
$$\text{s.t. } y_i = B\left(\beta l^{\rho_y} + (1-\beta)k^{\rho_y}\right)^{\frac{1}{\rho_y}}$$

Then the first order conditions are:

$$w = \eta\beta B^{\rho_y}\left(\frac{y_i}{l_i}\right)^{1-\rho_y} \qquad r = \eta(1-\beta)B^{\rho_y}\left(\frac{y_i}{k_i}\right)^{1-\rho_y}$$

Which we can rewrite as:

$$l_i = y_i\left(\frac{\eta\beta B^{\rho_y}}{w}\right)^{\frac{1}{1-\rho_y}} \qquad k_i = y_i\left(\frac{\eta(1-\beta)B^{\rho_y}}{r}\right)^{\frac{1}{1-\rho_y}}$$

Substituting into the production function and rearranging we can solve for the Lagrange multiplier:

$$\eta = \frac{1}{B}\left(\beta^{\sigma_y} w^{1-\sigma_y} + (1-\beta)^{\sigma_y} r^{1-\sigma_y}\right)^{\frac{1}{1-\sigma_y}}$$

where $\sigma_y = 1/(1-\rho_y)$. Note that because we set this up as a cost-minimization problem, this is the marginal cost. With CRS production, this is independent of the scale of production, which means that all firms have the same marginal cost whatever their size. And since large firms cannot drive down costs, there is no natural monopoly in this industry. This argument justifies the assumption of taking factor prices as given. It also justifies marginal cost pricing, so we can set the sales price $p_y = \eta$.

Constant marginal costs across firms also means that we have aggregation of factor demands. Since labor and capital are linear in output, we can simply sum up the factor demands to derive the aggregate factor demand curves:

$$L_y = \left(\frac{Y}{B}\right)\left(\frac{\beta p_y B}{w}\right)^{\sigma_y} \qquad K_y = \left(\frac{Y}{B}\right)\left(\frac{(1-\beta) p_y B}{r}\right)^{\sigma_y}$$

Where $Y = \sum_i y_i, L_y = \sum_i l_i, K_y = \sum_i k_i$. Thus, at any prices, we can treat aggregate output as if produced by a representative firm.

B.4 Production with a fixed factor

We now consider the same problem, but where capital is fixed at the firm level. The cost minimization problem:

$$\min_{l_i} wl_i + r\bar{k}_i$$
$$\text{s.t. } y_i = B\left(\beta l^{\rho_y} + (1-\beta)\bar{k}^{\rho_y}\right)^{\frac{1}{\rho_y}}$$

Then the first order condition is once more:

$$w = \eta \beta B^{\rho_y} \left(\frac{y_i}{l_i} \right)^{1-\rho_y}$$

Solving for l_i

$$l_i = y_i \left(\frac{\eta \beta B^{\rho_y}}{w} \right)^{\frac{1}{1-\rho_y}}$$

Substituting into the production function and rearranging we get that:

$$\eta = \left(\frac{w}{B} \right) \beta^{\frac{-\sigma_y}{\sigma_y - 1}} \left(1 - (1-\beta) \left(\frac{k_i B}{y} \right)^{\frac{\sigma_y - 1}{\sigma_y}} \right)^{\frac{1}{\sigma_y - 1}}$$

Which says that the marginal cost of producing output is just the product wage, w/B, adjusted for expenditure share. This adjustment itself if a function of the weight of labor in production (the β term) and an adjustment for the scarcity of capital.

It is relatively straightforward to show that the marginal cost is increasing in y:

$$\frac{d \log \eta}{d \log y} = \frac{1}{\sigma_y} \left(\frac{1}{\nu} - 1 \right) > 0$$

where: $\nu = 1 - (1-\beta) \left(\frac{k_i B}{y_i} \right)^{\frac{\sigma_y - 1}{\sigma_y}}$ which is bounded above by 1.

An increasing marginal cost curve means that both factor and product markets are competitive, so price-taking is a reasonable assumption (just as in the flexible factor case). Note also that if labor and capital are more substitutable then the marginal cost curve is less steep.

With competitive markets, firms choose output y_i so that the price p_y equals the marginal cost η. Thus, the capital-output ratio k_i/y_i is constant across firms, and is given by:

$$\frac{k_i}{y_i} = \frac{1}{B}\left((1-\beta)^{-1}\left(1-\beta^{\sigma_y}\left(\frac{Bp_y}{w}\right)^{\sigma_y-1}\right)\right)^{\frac{\sigma_y}{\sigma_y-1}}$$

That is, firms produce output in proportion to their initial level of capital.

Likewise, substituting back into the labor demand equation we see that firm-level labor demand is also linear in output:

$$l_i = y_i\left(\frac{p_y\beta B^{\rho_y}}{w}\right)^{\frac{1}{1-\rho_y}}$$

B.4.1 Aggregation

The forgoing equations mean that aggregation also holds in this case. That is, for any final goods price p_y and any market wage w, the total output $Y = \sum_i y_i$ from a large number of firms using this production technology is the same for any total capital $\bar{K}_y = \sum_i \bar{k}_i$ not matter the distribution of that capital across firms. In other words, the relationship between aggregate labor, capital, output, and prices is given by:

$$\bar{K}_y = \frac{Y}{B}\left((1-\beta)^{-1}\left(1-\beta^{\sigma_y}\left(\frac{Bp_y}{w}\right)^{\sigma_y-1}\right)\right)^{\frac{\sigma_y}{\sigma_y-1}}$$
$$L_y = \left(\frac{Y}{B}\right)\left(\frac{\beta p_y B}{w}\right)^{\sigma_y}$$

And so labor usage is given by:

$$L_y = \bar{K}_y(1-\beta)^{\frac{\sigma_y}{\sigma_y-1}}\left(\left(\frac{p_y B}{w}\right)^{1-\sigma_y}-\beta^{\sigma_y}\right)^{\frac{-\sigma_y}{\sigma_y-1}}$$

Given that k_i is fixed, there is nothing pinning down the return on capital, r. Moreover, any choice of r just selects a different level of profit for each firm, since the sum of profits and payments to capital always sum to $p_y y_i - w l_i$. However, if owners of firms and capital are the same, then the distinction between profits and capital payments is irrelevant. And so we can pick the equilbirium where r is the marginal product of capital, in which case profits are zero for all firms.

B.4.2 Connection to aggregate DRS technology

We now show how to approximate a CES production function with fixed capital with a DRS production function of the sort considered in Section B. That is, imagine we want to approximate the aggregate production function:

$$Y = B\left(\beta L_y^{\frac{\sigma_y - 1}{\sigma_y}} + (1-\beta)\bar{K}_y^{\frac{\sigma_y - 1}{\sigma_y}}\right)^{\frac{\sigma_y}{\sigma_y - 1}}$$

by choosing the values of $\tilde{\beta}$ and $\tilde{B}$ in:

$$Y = \tilde{B} L_y^{\tilde{\beta}}$$

Then for the approximation to be good near some fixed L_y^*, we require that the level and first (log) derivative of the approximating function match that of the true production function. That is:

$$\tilde{\beta} = \frac{\beta}{\beta + (1-\beta)\left(\frac{\bar{K}_y}{L_y^*}\right)^{\frac{\sigma_y - 1}{\sigma_y}}}$$

$$\tilde{B} = \frac{B\left(\beta L_y^{\frac{\sigma_y - 1}{\sigma_y}} + (1-\beta)\bar{K}_y^{\frac{\sigma_y - 1}{\sigma_y}}\right)^{\frac{\sigma_y}{\sigma_y - 1}}}{\left(L_y^*\right)^{\tilde{\beta}}}$$

PUBLICATIONS

Immigration and Local Inflation

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