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Globalization, Inflation Dynamics, and the Slope of the Phillips Curve

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Globalization, Inflation Dynamics, and the Slope of the Phillips Curve*

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Abstract

We study how international trade has affected U.S. inflation dynamics. We develop a multi-region New Keynesian model of an open monetary union in which regional Phillips curve slopes depend on exposure to imported goods. Guided by the model, we construct state-level measures of import penetration in final consumption and exploit cross-state variation to estimate regional Phillips curves over 1976-2017. Greater import exposure leads to significantly atter state-level Phillips curves. We estimate that rising import exposure accounts for roughly 40 percent of the flattening of the U.S. Phillips curve over this period. Increased import exposure also shifted the U.S. Phillips curve downward, reducing inflation by about 0.17 percentage point annually between 1976 and 2008. A reversal of globalization to early-1990s levels would substantially alter monetary policy trade-offs, reducing the sacrifice ratio by almost one-half. Our findings show that international trade has played a quantitatively important role in shaping U.S. inflation dynamics.

JEL Classification: E10, E30, F00, F40, F60

Keywords: Globalization, inflation, Phillips curve, international trade

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And insofar as globalization has affected the dynamics of inflation—through changes in the slope of the Phillips curve or the NAIRU (non-accelerating inflation rate of unemployment)—it may require some recalibration of policy responses.

Yellen (2006), *"Monetary Policy in a Global Environment"*

All in all, the conjecture that globalization has flattened the U.S. Phillips curve remains just that—a conjecture. More study is needed.

Obstfeld (2020), *"Global Dimensions of U.S. Monetary Policy"*

1 Introduction

The world economy has become substantially more globalized over the past fifty years. Alongside the marked increase in cross-border flows of goods, services, and financial assets, the United States has become increasingly integrated into the global economy. The expansion in trade in goods is particularly striking. As shown in Figure 1, U.S. imports of goods as a share of GDP more than tripled between 1970 and the early 2010s. This surge in imports has sparked an extensive research agenda on the wide-ranging economic consequences of globalization.[1] Yet despite this extensive body of work, how globalization—and trade in goods in particular—has shaped U.S. inflation dynamics remains an open question.

Understanding whether globalization has altered the behavior of inflation is central to both policymakers and academics. For policymakers, increased trade integration may affect price dynamics in ways that require adjustments in the conduct of monetary policy (Yellen, 2006). This concern arises even if globalization does not directly impair a central bank's ability to achieve its inflation objective.[2] Within academia, systematic attention to the implications of globalization for inflation and monetary policy emerged alongside increasing economic integration (Gali and Gertler, 2007). A central question in this literature is whether globalization has affected the relationship between unemployment and inflation. The fact that this relationship declined markedly over the same period in which U.S. import exposure increased, as illustrated in Figure 1, has led some researchers to suggest that greater import exposure may have flattened the U.S. Phillips curve.[3,4] Despite this suggestive evidence and theoretical work examining how trade integration may flatten the Phillips curve, the quantitative importance of globalization in explaining the diminished role of domestic slack in U.S. inflation remains unresolved (Obstfeld, 2020).[5]

[1] This research agenda includes the study of the reallocation of workers across industries and regions (Autor et al., 2013; Pierce and Schott, 2016; Caliendo et al., 2019; Greenland et al., 2019), labor market dynamics (Dix-Carneiro et al., 2023), and movements in the relative prices of imported goods (Amiti et al., 2020; Jaravel and Sager, 2019).

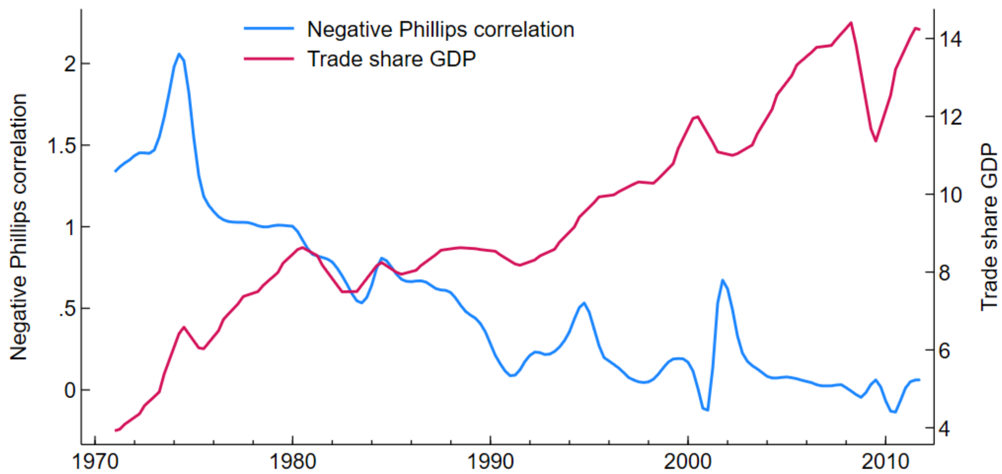
[2] Obstfeld (2020) discusses the channels through which global factors influence U.S. monetary policy tradeoffs.

[3] Figure 1 plots the Stock and Watson (2019) "Phillips correlation" derived from the accelerationist Phillips curve.

[4] Forbes (2019) argues that globalization may help explain the "Missing Inflation Puzzle"—the apparent weakening of the relationship between unemployment and inflation.

[5] See, for example, Sbordone (2007), Benigno and Faia (2016), and Guilloux-Nefussi (2020).

Figure 1: Stock-Watson "Phillips Correlation" and U.S. Goods Import Share



Notes: For Trade Share of GDP (red line), the figure plots the nominal value of total U.S. imports of goods as a share of U.S. nominal GDP as a percent on the right axis. The Negative Phillips correlation (blue line) is computed as the negative of the coefficient obtained from a rolling regression with a 33 quarter window starting in 1970 of the annual change in the 12-month core personal consumption expenditures inflation on the Congressional Budget Office unemployment gap.

This paper provides evidence that international trade affected inflation dynamics in the United States significantly by flattening the U.S. Phillips curve and shifting it downward since the late 1970s. To estimate the impact of import exposure on the U.S. Phillips curve, we propose an empirical methodology to estimate state-level Phillips curves based on the predictions of multi-region open-economy New Keynesian model in which regions differ in their exposure to imported goods. The model that we put forward delivers two different effects through which trade affects inflation in a region. First, a region's level of import exposure affects the sensitivity of inflation to movements in the unemployment rate—that is, the slope of the Phillips curve. Second, increases in a region's level of import exposure affect inflation directly by reflecting declines in import prices, thereby shifting the region's Phillips curve downward. The predictions of our model discipline our empirical strategy, which exploits cross-regional variation in inflation, unemployment, and import exposure to estimate U.S. regional Phillips curves.[6] We consider the period 1977–2017 to estimate these two effects. Our results provide evidence that more open U.S. states had flatter Phillips curves and that changes in import exposure shifted states' Phillips curves; albeit the evidence on the latter effect is less clear.[7] Relying on our estimates, we show that greater U.S. trade exposure contributed significantly to the flattening of the Phillips curve over the period considered. We explore additional counterfactual exercises shedding light on how monetary policy trade-offs—summarized by the U.S. "sacrifice ratio"—would change if the level of U.S. trade openness were to return to its early-1990s level.

[6] Previous work has focused on how supply-side factors related to openness affect inflation dynamics (Forbes, 2019). Our analysis takes these factors into account, but it also incorporates the effects on the slope of the Phillips curve.

[7] In line with recent research in both empirical macroeconomics and international trade, we rely on idiosyncratic regional variation for identification (Mian and Sufi, 2014; Chodorow-Reich, 2017; Nakamura and Steinsson, 2018; McLeay and Tenreyro, 2019; Autor et al., 2013). This strategy helps us overcome multiple issues that arise when one relies on variation along the time dimension only.

Our analysis proceeds in three parts. First, we develop a model of an open monetary union to characterize the mechanism linking trade openness and inflation dynamics. We extend a canonical New Keynesian small open economy (SOE) model (Galí and Monacelli, 2005) with staggered price adjustments to incorporate multiple regions that differ in their degree of trade openness vis-à-vis the rest of the world. In the baseline specification, regions within the monetary union trade final goods with the rest of the world but not with one another—an "islands" structure.[8]

The model's equilibrium conditions yield regional Phillips curves for final consumer price inflation that differ from those in closed-economy settings along two dimensions. First, the standard domestic determinants of inflation—the unemployment gap and expected inflation—are augmented by changes in import penetration, defined as the share of total regional final expenditure devoted to foreign-produced goods. Second, the slope of a region's Phillips curve—that is, the sensitivity of inflation to deviations of unemployment from its long-run level—depends on the region's steady-state level of import penetration. We refer to the first dimension as the *direct effect* of trade openness on inflation, as it captures the level effect arising from lower import prices associated with reduced trade barriers, higher foreign productivity, or other global factors.[9] We label the second dimension the *indirect effect*, as trade openness influences inflation dynamics by altering the responsiveness of real marginal costs—and thus prices—to demand-driven fluctuations in economic activity. Under standard parametric assumptions, real marginal costs respond less to demand shocks in more open regions that consume a larger share of imported goods. Thus, more closed regions have steeper region-specific Phillips curves.

Taken together, the model implies that, conditional on identical unemployment gaps and inflation expectations, cross-regional differences in both contemporaneous changes and levels of import penetration generate systematic differences in inflation dynamics across regions. Hence, to the extent that there is heterogeneity in regions' exposure to trade in the United States, we should observe different inflation dynamics across these regions.

In the second part of the analysis, we quantify these effects empirically over the period 1976–2016. Guided by the model's predictions, we exploit cross-state variation in trade exposure, inflation, and unemployment to estimate regional Phillips curves. The use of regional data provides at least two identification advantages.[10] First, aggregate shocks can be absorbed through time fixed effects. This feature allows us to address key challenges in traditional Phillips curve estimation by mitigating the endogeneity bias arising from the systematic response of monetary policy to cost-push shocks (McLeay and Tenreyro, 2019; Fitzgerald et al., 2020). It also enables us to control for time variation in aggregate inflation expectations and potential output (Hazell et al., 2022). To further address potential issues of endogeneity bias due to movements in unemployment being correlated with trade exposure, we also consider a variable instrument for unemployment based on Nekarda and Ramey

[8] We also develop extensions that allow for trade in intermediate inputs and interregional trade. We abstract from these features in the empirical analysis due to data limitations discussed in the paper.

[9] Global forces such as exchange rate movements, foreign inflation, commodity price fluctuations, and supply chain disruptions affect domestic inflation primarily through import prices.

[10] See Furlanetto and Lepetit (2025) for a detailed discussion on the advantages of approaches that do not rely on aggregate data to estimate the Phillips curve.

(2011) that accounts for this potential bias.[11] Overall, our empirical strategy integrates recent methodological advances from both empirical macroeconomics and international trade within a unified and parsimonious framework.

A second advantage of using regional data is that cross-sectional variation generates substantially more dispersion in trade exposure than aggregate time-series data. A large literature in international trade leverages regional heterogeneity in exposure to trade shocks to identify causal effects (Autor et al., 2013; Kovak, 2013). We implement a shift-share strategy—allowing us to address potential endogeneity in trade exposure—to construct a state-level measure of import penetration in final consumption goods that aligns with the concept of openness in our model.[12] We begin by constructing a national measure of import penetration at the NAICS four-digit level for sectors in agriculture, mining, and manufacturing in each year. We then combine these national import penetration rates with state-level expenditure shares at the detailed product level (e.g., automobiles, apparel, jewelry, furniture) using 1994 data from the Bureau of Labor Statistics Consumer Expenditure Survey public-use microdata. This procedure yields a time-varying, state-specific measure of the share of final consumption sourced from abroad, ranging from zero (all consumption domestically produced) to one (all consumption imported).

Figure 2 illustrates the cross-state distribution of import penetration in 1977 and 2016. The median state had an import penetration rate of roughly 3 percent in 1977, rising to approximately 12 percent by 2016. Importantly, states differ meaningfully in their exposure to this increase in trade openness, providing the variation that underpins our empirical estimates.

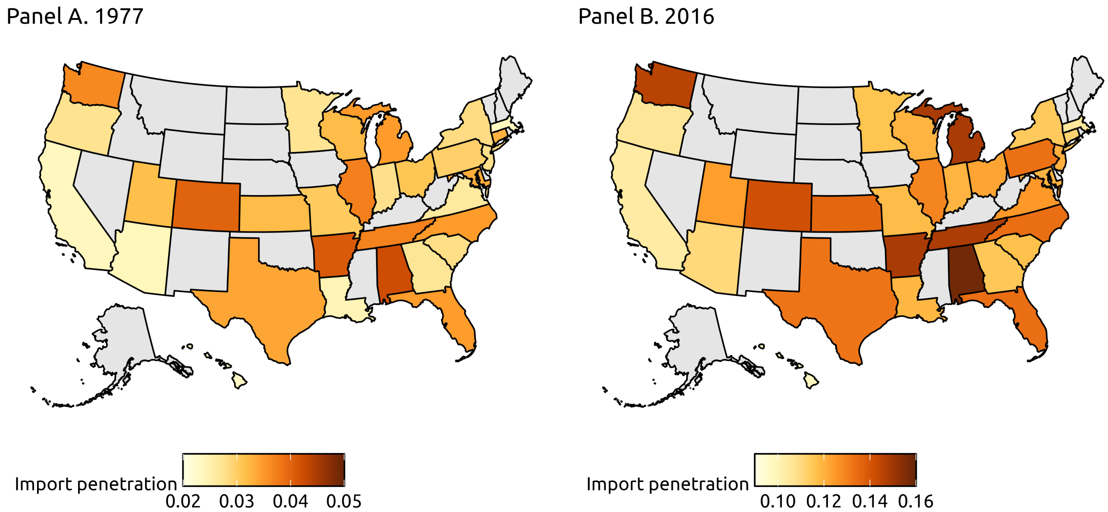
Our estimates provide evidence that the slopes of state-level Phillips curves for non-housing inflation differ systematically across states, consistent with the presence of the *indirect effect* of trade openness. In particular, we find that more open states exhibit flatter Phillips curves, in line with reasonable parametric assumptions of our theoretical model. To assess the economic significance of our estimates, consider two regions in a given year: one relatively open, with an import exposure of 0.12, and one more closed, with exposure of 0.08. According to our instrumental variables (IV) estimates, the slope of the Phillips curve in the more open region is 0.02, compared with 0.07 in the more closed region—more than three times larger. By contrast, we find no statistically significant evidence of an *indirect effect* of trade for housing inflation, consistent with housing services being largely nontradable.

We construct the aggregate U.S. Phillips curve by combining the estimated regional Phillips curves for both non-housing and housing components and aggregating regional Phillips curves across states. Using our estimates and the national measure of import exposure, we find that the slope of the U.S. Phillips curve declined from 0.13 in 1977 to 0.07 in 2017. These magnitudes are consistent with Hazell et al. (2022). Importantly, our framework allows us to isolate the portion of this decline attributable to rising trade openness. Overall, the results indicate that globalization has had a quantitatively meaningful effect on U.S. inflation dynamics.

[11] More generally, this type of identification issues arise from the fact that the Phillips curve relationship holds conditional on demand shocks.

[12] This approach also allows us to circumvent the absence of state-level trade data prior to the early 2000s.

Figure 2: Import penetration across U.S. states: 1977 and 2016



Notes: Import penetration in final consumption constructed following the procedure described in Section 4.2. Data sources include the Consumer Expenditure Survey (CE) public-use microdata provided by the Bureau of Labor Statistics (BLS), the NBER-CES Manufacturing Database, the Bureau of Economic Analysis (BEA) detailed GDP-by-Industry data, the U.S. Census Bureau, and the United States Trade Representative (USITC).

Turning to the direct effect of trade openness, our estimates are consistent with the theoretical framework: states experiencing larger contemporaneous increases in import penetration exhibit lower non-housing inflation. The magnitude of these estimates aligns with cross-sectional evidence on trade elasticities in the existing literature (Simonovska and Waugh, 2014; Caliendo and Parro, 2015). However, the estimates are imprecisely measured and do not attain conventional levels of statistical significance. We hypothesize that this result may reflect the high cross-state correlation in changes in import penetration, which limits the independent variation available to precisely estimate the direct effect.

In additional counterfactual exercises, we find three main results. First, the substantial increase in U.S. trade exposure over the past four decades contributed meaningfully to the flattening of both state-level Phillips curves and the aggregate U.S. Phillips curve. We estimate that approximately 40 percent of the observed decline in the slope of the national Phillips curve can be attributed to increased import exposure. Second, absent the trade-induced flattening, annual inflation would have been roughly 0.5 percentage points lower at the height of the Great Recession and about 0.3 percentage points higher in the years immediately preceding the COVID-19 pandemic. Third, a sizable reversal of globalization—were it to occur—would constitute a structural shift in inflation dynamics that monetary policy would need to account for. In a deglobalization scenario in which the average U.S. state-level import penetration measure declines from 12.3 to 6.5 percent—the levels observed for 2017 and 1991, respectively—the sacrifice ratio—that is, the increase in the unemployment rate gap required to lower inflation 1 percentage point—would decline from about 16.7 percent previously to about 11.1 percent with lower import penetration. Moreover, if we assume that this decline occurs over a period of eight years, the direct effect of this retrench in globalization boosts U.S. headline CPI inflation by about 0.5 percentage points per year over this period.

Related Literature Our work is related to several strands of the literature in macroeconomics and international trade that study the interaction between globalization and inflation dynamics. In particular, our work relates to (i) the literature on global determinants of inflation (Borio and Filardo, 2007; Ciccarelli and Mojon, 2010; Bianchi and Civelli, 2015; Forbes, 2019; Obstfeld, 2020), (ii) open-economy New Keynesian models of inflation (Clarida et al., 2002; Galí and Monacelli, 2005; Sbordone, 2007; Benigno and Faia, 2016), (iii) empirical work on trade and prices (Amiti et al., 2014, 2019; Jaravel and Sager, 2019), and (iv) studies exploiting regional variation to identify macroeconomic relationships (Charles et al., 2018; Nakamura and Steinsson, 2018; Hazell et al., 2022). Our paper is most closely related to the studies that have focused on using cross-regional variation to identify regional Phillips curves (Kumar and M. Orrenius, 2016; Fitzgerald et al., 2020; Hazell et al., 2022). Recent work analyzes the spatial dimension of monetary policy shocks (Pedemonte and Herreño, 2022).

A large body of work has examined whether globalization has altered the inflation process by increasing the importance of global factors relative to domestic slack. Early contributions by Borio and Filardo (2007) and Ciccarelli and Mojon (2010) provide cross-country evidence suggesting that global economic conditions play a significant role in shaping domestic inflation dynamics. More recent work by Eickmeier and Ng (2015) and Kabukçuoğlu and Martínez-García (2018) further emphasizes the role of global slack and international comovement in inflation.

This literature is closely related to the question of whether globalization has contributed to the flattening of the Phillips curve. Forbes (2019) argues that global factors—such as import competition, global value chains, and exchange rate movements—may help explain the weakening relationship between domestic slack and inflation, often referred to as the "missing inflation" puzzle. Similarly, Obstfeld (2020) highlights the growing importance of global forces for monetary policy and the challenges they pose for interpreting domestic inflation dynamics.

Our paper complements this literature by providing a micro-founded and empirically tractable framework that links global integration to inflation dynamics through trade exposure. Rather than relying on aggregate cross-country correlations, we exploit cross-state variation within the United States to identify how differences in exposure to international trade affect inflation dynamics. In doing so, we provide direct evidence that globalization reduces the sensitivity of inflation to domestic slack.

Our theoretical framework builds on the open-economy New Keynesian literature that studies inflation dynamics in the presence of international trade. Seminal contributions by Clarida et al. (2002) and Galí and Monacelli (2005) show that openness affects inflation through its impact on marginal costs and the terms of trade. In these models, trade integration alters the transmission of domestic shocks to inflation by changing how wages and prices adjust in equilibrium.

Subsequent work has further explored the implications of globalization for inflation dynamics.

Sbordone (2007) emphasizes that increased competition from foreign producers may reduce firms' pricing power, thereby weakening the response of inflation to domestic economic activity. Benigno and Faia (2016) highlight the role of international trade in shaping pass-through and inflation persistence in open economies.

Our paper extends this literature along two important dimensions. First, we introduce heterogeneity in trade openness across regions within a monetary union. This feature allows us to generate cross-sectional variation in Phillips curve slopes, providing a novel empirical implication: more open regions should exhibit flatter Phillips curves. Second, we show that trade openness affects inflation through two distinct channels: a direct effect operating through import prices and an indirect effect operating through the sensitivity of marginal costs to domestic slack. These mechanisms allow us to map theoretical predictions directly into an empirically implementable specification, bridging a gap between theory and data in the open-economy Phillips curve literature.

A related strand of the literature examines how international trade affects domestic prices and inflation through changes in import competition and input costs. Amiti et al. (2014) document that firm-level pricing behavior and exchange rate pass-through are shaped by firms' participation in international trade. Jaravel and Sager (2019) provide evidence that increased import competition—particularly from China—has led to substantial declines in U.S. consumer prices.

These studies provide micro-level evidence for the direct effect of trade on inflation emphasized in our framework. In our model, increases in import penetration—driven by lower import prices or reduced trade costs—directly lower consumer price inflation by shifting expenditure toward cheaper foreign goods. While our empirical estimates of this direct effect are imprecise, they are consistent in sign with the findings in this literature. Importantly, our approach complements these studies by embedding these price effects within a general equilibrium framework that also captures how trade affects the slope of the Phillips curve.

Our empirical strategy is closely related to the literature that exploits regional variation in exposure to trade shocks to identify causal effects. Autor et al. (2013) demonstrate how differential exposure to Chinese import competition across U.S. regions can be used to study labor market outcomes. McLeay and Tenreyro (2019) and Hazell et al. (2022).

We build on this literature by applying similar identification strategies to study inflation dynamics. Specifically, we construct state-level measures of import penetration using a shift-share approach and exploit cross-state variation to estimate regional Phillips curves. This approach allows us to overcome key identification challenges present in aggregate time-series analyses, such as the endogeneity of monetary policy and the limited variation in aggregate trade exposure.

By bringing together methods from international trade and empirical macroeconomics, our paper provides new evidence on how globalization affects inflation dynamics within a country. To our knowledge, this is the first paper to use regional variation in trade exposure to quantify the impact of globalization on the slope of the Phillips curve.

Taken together, our findings contribute to the literature by providing a unified framework that integrates insights from open-economy macroeconomics and international trade. We show that globalization affects inflation dynamics through both level and slope effects, and we provide empirical evidence that these effects are quantitatively important for understanding the evolution of the U.S. Phillips curve. Our results offer a structural explanation for the weakening relationship between inflation and domestic slack and highlight the importance of considering trade exposure when evaluating monetary policy trade-offs.

Roadmap The remainder of this paper is organized as follows. Section 2 lays out our model. Section 3 shows how we use the model to discipline our empirical framework and discusses the appealing features of our approach to study the effects of openness on the Phillips Curve. Section 4 presents our data and empirical analysis. Section 5 conducts counterfactual experiments relying on our previous estimates. Section 6 presents a series of robustness exercises. Section 7 concludes.

2 A Multi-region Open Economy New-Keynesian Model

In this section we develop a model of the U.S. economy with sticky prices. Our framework extends a benchmark open economy model with nominal rigidities à la Calvo (Galí and Monacelli, 2005) to multiple regions that differ in their degree of openness towards the rest of the world. The United States is assumed to be an open economy composed of multiple regions indexed by $r \in \{1, \dots, I\}$, each of which is a small open economy relative to the rest of the world. We denote by ν_{rt} the share of total U.S. population located in region r at time t, such that $\sum_{r=1}^{I} \nu_{rt} = 1$ for every t. Our baseline model does not incorporate internal trade across regions in the United States. Thus, we can think of these regions as islands. Appendix D shows how to extend the model to include internal trade and discusses the implications for our empirical analysis in Section 4.

2.1 Households

Region r is inhabited by a representative household that seeks to maximize the expected value of its lifetime utility,

$$E_0 \sum_{t=0}^{\infty} \beta^t U\left(C_{rt}, N_{rt}; Z_{rt}\right), \tag{1}$$

where C_{rt} is per capita consumption of a composite consumption index in region r, N_{rt} is per capita employment in region r, and Z_{rt} is an exogenous stochastic preference shifter in region r. The composite consumption index is defined by

$$C_{rt} = \left((1 - \upsilon_r)^{\frac{1}{\eta}} \left(C_{H,rt}\right)^{1-\frac{1}{\eta}} + (\upsilon_r)^{\frac{1}{\eta}} \left(C_{F,rt}\right)^{1-\frac{1}{\eta}} \right)^{\frac{\eta}{\eta-1}}, \tag{2}$$

where $C_{H,rt}$ is an index of consumption of goods produced locally—that is, produced in region r—which is given by

$$C_{H,rt} \equiv \left(\int_0^1 C_{H,rt}\left(i\right)^{\frac{\varepsilon-1}{\varepsilon}} di \right)^{\frac{\varepsilon}{\varepsilon-1}}, \tag{3}$$

where $i \in [0,1]$ indexes varieties and $C_{F,rt}$ is the index of consumption of goods produced abroad.[13] We interpret the parameter $\upsilon_r \in [0,1]$ as a measure of openness that can differ across regions. This is in line with the evidence shown in Figure 2. More open regions will have higher values of υ_r, and therefore lower home bias.[14] The parameter $\eta \geq 1$ determines the trade elasticity, which in this model is given by $\eta - 1 \geq 0$.

Households maximize utility in (1) subject to the sequence of budget constraints given by

$$\int_0^1 P_{H,rt}\left(i\right) C_{H,rt}\left(i\right) di + P_{F,rt}C_{F,rt} + E_t\left[Q_{t,t+1}D_{rt+1}\right] \leq D_{rt} + W_{rt}N_{rt}$$

for $t = 0, 1, \ldots$ where D_{rt+1} is the nominal payoff in period $t+1$ of the portfolio held at the end of period t.[15] We assume that there are complete international financial markets and that households have access to a complete set of contingent claims that are traded internationally and frictionless.

In what follows, we assume that the period utility function $U\left(C_{rt}, N_{rt}; Z_{rt}\right)$ takes the form

$$U\left(C_{rt}, N_{rt}; Z_{rt}\right) = Z_{rt}\left(\frac{C_{rt}^{1-\sigma} - 1}{1-\sigma} - \frac{N_{rt}^{1+\varphi}}{1+\varphi}\right) \tag{4}$$

where $\sigma \geq 0$ and $\varphi \geq 0$ determine the curvature of the utility of consumption and the disutility of labor, respectively.

2.2 Firms

2.2.1 Technology

There is a continuum of firms in each region, each producing a tradable variety. Varieties can be costlessly traded with the rest of the world, but we assume that these cannot be traded across regions. As in the previous section, a firm producing variety i is indexed by $i \in [0,1]$. We assume that a typical firm in region r has technology

$$Y_{rt}\left(i\right) = A_{rt}N_{rt}\left(i\right), \tag{5}$$

where A_{rt} is an exogenous and stochastic productivity shock affecting region r.

[13] As shown in appendix D, the model can be extended to allow for trade across regions, in which case $C_{H,rt}$ would be replaced by $C_{L,rt}$, an index of goods produced in any region r in the monetary union.

[14] Alternatively, we could have assumed that trade openness was parameterized by iceberg-type trade costs. We choose υ_r as the relevant parameter following Galí and Monacelli (2005).

[15] The portfolio can include shares in local firms.

We assume that firms set prices in a staggered fashion, as in Calvo (1983). Hence, a measure $1-\theta$ of firms set new prices each period, with an individual firm's probability of re-optimizing in any given period being independent of the time elapsed since it last reset its price. Firms can sell their goods locally and also export them to the rest of the world. In line with Galí and Monacelli (2005), we assume that these firms engage in producer currency pricing when choosing the price at which they offer their goods.

2.3 Exports

In addition to domestic households demanding goods in region r, there is also a demand by the rest of the world for each variety $i \in [0,1]$ produced in each region. We assume that the demand for exports of good $i \in [0,1]$ produced in region r is given by

$$X_{rt}(i) = \left(\frac{P_{H,rt}(i)}{P_{H,rt}} \right)^{-\varepsilon} X_{rt}, \tag{6}$$

where $X_{rt} \equiv \left(\int X_{rt}(i)^{\frac{\varepsilon-1}{\varepsilon}} di \right)^{\frac{\varepsilon}{\varepsilon-1}}$ is an index across varieties, summarizing aggregate exports from region r to the rest of the world. Turning to aggregate exports, we assume that they are given by

$$X_{rt} = \frac{\upsilon^*}{I} \left(\frac{P_{H,rt}}{\mathcal{E}_t P_t^*} \right)^{-\eta} Y_t^*, \tag{7}$$

where, similar to the case of a region of the U.S. economy, $\upsilon^* \in [0,1]$ parameterize the degree of openness by the rest of the world vis-à-vis the United States, I is the number of regions that form the U.S. economy, $\mathcal{E}_t$ is the nominal exchange rate expressed in terms of U.S. dollars per rest of the world currency units, P_t^* denotes the price of goods produced abroad expressed in local currency units, and Y_t^* is real GDP in the rest of the world. As will become clearer in the next section, the previous expression implicitly assumes that preferences of households in the rest of the world are very similar to those of domestic households.

2.4 Government Policy

The monetary authority conducts a common policy for all regions $r = 1, \ldots, I$. We assume that the policy takes the form of the following interest rate rule

$$\hat{r}_t = \phi_\pi (\pi_t - \bar{\pi}_t) - \phi_u (\hat{u}_t - \bar{u}_t) \tag{8}$$

where hats denote deviations from the zero-inflation steady state, and $\bar{\pi}_t$ and $\bar{u}_t$ denote aggregate inflation and unemployment targets by the central bank. We assume that the parameters determining the rule are such that the economy has a unique locally bounded equilibrium. Aggregate inflation and unemployment are in turn defined as population-weighted averages across regions:

$$\pi_t = \sum_{r=1}^{I} \nu_{rt}\pi_{rt} \quad \text{and} \quad \hat{u}_t = \sum_{r=1}^{I} \nu_{rt}\hat{u}_{rt}.$$

We consider a passive fiscal authority that does not spend, issues debt, nor levies taxes.

2.5 Optimality Conditions

2.5.1 Households

From the solution of the households' maximization problem, we obtain that the optimal allocation of expenditures across locally produced varieties is characterized by the conditional demand function

$$C_{H,rt}(i) = \left(\frac{P_{H,rt}(i)}{P_{H,rt}}\right)^{-\varepsilon} C_{H,rt}, \tag{9}$$

where the price index $P_{H,rt}$ is given by

$$P_{H,rt} = \left(\int_0^1 P_{H,rt}(i)^{1-\varepsilon}\, di\right)^{\frac{1}{1-\varepsilon}}, \tag{10}$$

and $C_{H,rt}$ denotes the consumption index of locally produced goods. These prices and quantities are such that $\int_0^1 P_{H,rt}(i)\, C_{H,rt}(i) = P_{H,rt}C_{H,rt}$.

The index $C_{H,rt}$, in turn, as well as the non-local final consumption index are determined by

$$C_{H,rt} = (1-\upsilon_r)\left(\frac{P_{H,rt}}{P_{rt}}\right)^{-\eta} C_{rt} \quad \text{and} \quad C_{F,rt} = \upsilon_r \left(\frac{P_{F,rt}}{P_{rt}}\right)^{-\eta} C_{rt}, \tag{11}$$

where the price index of the final consumption composite basket is

$$P_{rt} = \left[(1-\upsilon_r)\,(P_{H,rt})^{1-\eta} + \upsilon_r\,(P_{F,rt})^{1-\eta}\right]^{\frac{1}{1-\eta}} \tag{12}$$

for the final consumption good in region r.

Given the functional form assumed in (4), we can rewrite the intratemporal optimality condition by the households as

$$C_{rt}^{\sigma} N_{rt}^{\varphi} = \frac{W_{rt}}{P_{rt}}, \tag{13}$$

and the intertemporal optimality conditions as

$$\beta\left(\frac{C_{rt+1}}{C_{rt}}\right)^{-\sigma}\left(\frac{P_{rt}}{P_{rt+1}}\right) = Q_{t,t+1}. \tag{14}$$

These set of conditions can be rewritten to deliver the more usual Euler equation

$$\beta R_t \mathbb{E}_t \left\{ \left(\frac{C_{rt+1}}{C_{rt}} \right)^{-\sigma} \left(\frac{P_{rt}}{P_{rt+1}} \right) \right\} = 1 \tag{15}$$

where $R_t = 1/\mathbb{E}_t\left[Q_{t,t+1}\right]$ denotes the gross return on a riskless one-period bond.

2.5.2 Price Setting by Firms

Let $Y_{rt}(i)$ denote total output by firm i located in region r at time t. This firm's nominal profits at time t are given by

$$\Pi_{rt}(i) = P_{H,rt}(i) Y_{rt}(i) - W_t N_{rt}(i). \tag{16}$$

Given that firms face sticky prices and they can only adjust their price in period t with probability $1-\theta$, if firm i is able to update its price, $P_{H,rt}(i)$, in period t, it will do so to maximize the objective

$$\sum_{k=0}^{\infty} \theta^k \mathbb{E}_t \left\{ Q_{t,t+k} \left(P_{H,rt}\left(i\right) Y_{rt+k}\left(i\right) - W_{rt+k} N_{rt}\left(i\right) \right) \right\} \tag{17}$$

subject to firm i's output satisfying demand given by

$$Y_{rt}\left(i\right) = C_{H,rt}\left(i\right) + X_{rt}\left(i\right) = \left(\frac{P_{H,rt}\left(i\right)}{P_{H,rt}} \right)^{-\varepsilon} \left(C_{H,rt} + X_{rt} \right) = \left(\frac{P_{H,rt}\left(i\right)}{P_{H,rt}} \right)^{-\varepsilon} Y_{H,rt} \tag{18}$$

where the second equality follows from equations (6) and (9) and by defining $Y_{H,rt} \equiv C_{H,rt} + X_{rt}$. Let $MC_{rt} \equiv \frac{W_{rt}}{P_{H,rt} A_{rt}}$ denote the real marginal cost faced by all firms in terms of domestic goods produced in region r at time t. Solving the problem of an updating firm implies that prices will be chosen such that

$$P_{H,rt}\left(i\right) = \frac{\varepsilon}{\varepsilon - 1} \frac{\sum_{k=0}^{\infty} \left(\theta\beta\right)^k \mathbb{E}_t \left\{ \frac{U'(C_{rt+k})}{P_{rt+k}} \left(MC_{rt+k} P_{H,rt+k}^{1+\varepsilon} Y_{H,rt+k} \right) \right\}}{\sum_{k=0}^{\infty} \left(\theta\beta\right)^k \mathbb{E}_t \left\{ \frac{U'(C_{rt+k})}{P_{rt+k}} P_{H,rt+k}^{\varepsilon} Y_{H,rt+k} \right\}}. \tag{19}$$

Given that the right-hand side (19) does not depend on i, all firms resetting prices at time t will choose the same price.

2.6 Equilibrium

In Appendix A we show that a log-linear approximation of the model around a zero inflation steady state delivers a relationship between domestic price inflation, $\pi_{H,rt}$, and real marginal cost given by

$$\pi_{H,rt} = \kappa \widetilde{mc}_{rt} + \beta \mathbb{E}_t \left[\pi_{H,rt+1} \right], \tag{20}$$

where domestic price inflation is defined by $\pi_{H,rt} \equiv p_{H,rt} - p_{H,rt-1}$ where $p_{H,rt} \equiv \log P_{H,rt}$, and $\widetilde{mc}_{rt} = mc_{rt} - mc_r$ where $mc_{rt} = \log MC_{rt}$ and $mc_r = -\log \frac{\varepsilon}{\varepsilon - 1}$ is the log of the real marginal cost in the zero inflation steady state. In (20), κ is a structural parameter given by

$$\kappa \equiv \frac{(1-\theta\beta)(1-\theta)}{\theta}. \tag{21}$$

Note that this parameter only depends on the probability of a firm readjusting its price in any given period, θ, and households' discount factor, β. According to equation (20), the pass-through from real marginal cost to domestic price inflation does not vary across regions. In particular, the degree of trade openness—a key source of regional heterogeneity in the model—does not enter this relationship and the relation between domestic price inflation and the real marginal cost is not affected by a region's degree of openness. However, as the next lemma shows, trade openness comes into play once we express the real marginal cost in terms of unemployment because trade affects the elasticity of real marginal cost with respect to the unemployment gap.

Lemma 1. *In a log-linear approximation of the model around its zero inflation steady state, domestic price inflation in region r can be written in terms of unemployment gaps and expected domestic price inflation as*

$$\pi_{H,rt} = -\kappa_r \widehat{u}_{rt} + \beta \mathbb{E}_t \left[\pi_{H,rt+1}\right], \tag{22}$$

where $\widehat{u}_{rt} = u_{rt} - u^n_{rt}$ denotes deviations of region r's unemployment rate, defined by $u_{rt} = \log U_{rt}$ where $U_{rt} = 1 - N_{rt}$, from its equilibrium level of unemployment in the absence of nominal rigidities and conditional on y^*_t. The region-specific coefficient, κ_r, is given by $\kappa_r \equiv \kappa(\varphi + \sigma^\upsilon_r)$ and

$$\sigma^\upsilon_r \equiv \frac{\sigma}{(1-\upsilon_r) + \upsilon_r \left[\sigma\eta + (1-\upsilon_r)(\sigma\eta - 1)\frac{1-\phi_r}{\upsilon_r}\right]}, \tag{23}$$

where $\phi_r \equiv \frac{1-\upsilon_r}{(1-\upsilon_r)+\frac{\upsilon^*}{I}}$.

Proof. See Appendix (B) □

Equation (22), which we refer to as the domestic price inflation Phillips curve, shows that the pass-through from regional unemployment gaps into domestic price inflation, κ_r, can vary across regions. This relationship emerges because the equilibrium relationship between marginal cost and unemployment is shaped by adjustments in relative prices that depend on substitution and income effects dependent on how open each region is. For instance, in a closed economy, $\upsilon_r \to 0$ for all r, and $\kappa_r = \kappa(\varphi + \sigma)$—as in the standard closed-economy New Keynesian model—and does not depend on the elasticities σ and η. If we consider the case of a single region and $\upsilon_r = \frac{\upsilon^*}{I}$, then $1 - \phi_r = \upsilon_r$ and the model boils down to the one developed by Galí and Monacelli (2005).

In this model, regional Phillips curves will be flatter for more open regions—that is, regions with higher values of υ_r—as long as $\sigma\eta > 1$.[16] Under this parameter restriction, more open regions experience a smaller adjustment in their terms of trade—defined as $S_{rt} \equiv \frac{P_{F,rt}}{P_{H,rt}}$—and therefore wages, necessary to absorb a change in domestic output (relative to world output).

To develop the intuition underlying this result, consider the extreme case of a closed region. If output produced in this region increases, the region has more resources to spend, but it only spends these on domestically produced goods. Hence, export prices, and therefore wages, have to drop in order for the foreign economy to buy part of this additional output. If the region is more open, then part of this additional expenditure is on goods produced abroad. Hence, the foreign economy also experiences an increase in its income leading to increased demand for the good produced in the domestic region implying that wages do not have to drop by as much. Hence, more openness dampens the impact of that adjustment on marginal cost and inflation. This effect is pointed out by Clarida et al. (2002) and Galí and Monacelli (2005).[17]

We turn now to the relationship between domestic and CPI inflation with the aim of deriving an expression for regional Phillips curves for final consumption price inflation in region r, $\pi_{rt} \equiv \frac{P_{rt}}{P_{rt-1}} - 1$. Consider the share of total final consumption expenditure by region r spent on goods produced abroad:

$$\Lambda_{rt} \equiv \frac{P_{F,rt} C_{F,rt}}{P_{rt} C_{rt}}. \tag{24}$$

The share $1 - \Lambda_{rt}$ provides a measure of home bias and a sufficient statistic for the welfare gains from trade openness in a large class of neoclassical models of international trade (Arkolakis et al., 2012). The literature on the welfare gains from trade has emphasized the relevance of this result because this share can be constructed using aggregate expenditure and production data without relying on price data.

It turns out that we can rely on changes in import penetration rather than import prices to obtain a Phillips curve for final price inflation that incorporates the effects of changes in the prices of goods brought from abroad. The following lemma formalizes this result.

Lemma 2. *Final consumption price inflation in region r can be expressed approximately in terms of domestic price inflation and changes in import penetration as*

$$\pi_{rt} = \pi_{H,rt} - \frac{\Delta \Lambda_{rt}}{\eta - 1}, \tag{25}$$

where $\Delta \Lambda_{rt} = \Lambda_{rt} - \Lambda_{rt-1}$ denotes changes in import penetration over time.

Proof. Equation (12) implies that $(1 + \pi_{rt})^{1-\eta} = (1 + \pi_{H,rt})^{1-\eta} \left(\frac{\left(\frac{1-\upsilon_r}{\upsilon_r}\right) + S_{rt}^{1-\eta}}{\left(\frac{1-\upsilon_r}{\upsilon_r}\right) + S_{rt-1}^{1-\eta}} \right)$ and that S_{rt} and Λ_{rt} are such that $S_{rt}^{\eta-1} = \frac{\upsilon_r}{1-\upsilon_r} \frac{1-\Lambda_{rt}}{\Lambda_{rt}}$. Hence, substituting the latter condition into the former implies

[16] We show in Appendix B that under $\sigma\eta > 1$, $\frac{\partial \sigma_r^\upsilon}{\partial \upsilon_r} < 0$. For the case in which $\upsilon_r = \upsilon^*/I$ (Galí and Monacelli, 2005), it can be easily seen that $\frac{\partial \sigma_r^\upsilon}{\partial \upsilon_r} < 0$ boils down to a simple expression.

[17] Similar intuition applies to the extensions of our baseline model to intermediate inputs and regional trade presented in Appendices C and D, respectively.

that

$$(1+\pi_{rt})^{1-\eta} = (1+\pi_{H,rt})^{1-\eta}\left(\frac{1-\Lambda_{rt}}{1-\Lambda_{rt-1}}\right)^{-1},$$

which delivers $\pi_{rt} \approx \pi_{H,rt} - \frac{\Delta\Lambda_{rt}}{\eta-1}$. $\square$

Lemma 2 implies that, given inflation of goods produced in region r, changes in region r's import penetration and an estimate of the trade elasticity, $\eta - 1$, are sufficient to obtain total inflation in the region. Therefore, we do not require data on prices or inflation of goods produced outside the U.S.. This feature of the model is very convenient for the empirical methodology that we follow in Section 4, precisely because data on import and export prices are not available at the regional level in the U.S.. Instead, we will exploit the fact that (24) can be restated as

$$\Lambda_{rt} = \frac{IM_{rt}}{VA_{rt} - (EX_{rt} - IM_{rt})}, \tag{26}$$

where $IM_{rt} \equiv P_{F,rt}C_{F,rt}$ denote imports, $EX_{rt} \equiv P_{H,rt}X_{rt}$ exports, and $VA_{rt} \equiv P_{H,rt}Y_{H,rt}$ value added in region r. The international trade literature has exploited the fact that all these variables are readily available at the national level for many countries. However, obtaining these measures or proxies of them at the regional level for the U.S. carries particular challenges that we address and discuss in Section 4.

We will refer to the statistic Λ_{rt} as import penetration from now on. Note that in the model, equation (9) implies that import penetration is given by $\Lambda_{rt} = \upsilon_r \left(\frac{P_{F,rt}}{P_{rt}}\right)^{1-\eta}$, which in a zero inflation steady state of the model with $P_{F,rt} = P_{rt}$ becomes $\Lambda_r = \upsilon_r$. This result clarifies the tight relationship between υ_r and the degree of openness of region r.

We can now obtain the expression for the CPI Phillips curve of region r. Relying on lemmas 1 and 2, we obtain regional final consumption price inflation, which depends on regional unemployment—a measure of economic activity—and changes in import penetration, which summarizes changes in import prices.

Proposition. *Final consumption price inflation in region r can be expressed in terms of regional unemployment, expected domestic foreign inflation, and changes in import penetration for region r as*

$$\pi_{rt} = -\kappa_r \widehat{u}_{rt} + \beta \mathbb{E}_t\left[\pi_{H,rt+1}\right] - \frac{\Delta\Lambda_{rt}}{\eta - 1} \tag{27}$$

where κ_r is as in Lemma (1).

Note that our regional Phillips curves for final consumption price inflation depend on expected future domestic price inflation. We could rewrite (27) in terms of expected future final consumption price inflation and changes in terms of trade as we first did in the proof of Lemma 2. However, a previously pointed out, we do not want to do this because the empirical analysis that we carry out in the remainder of the paper exploits the fact that we can compute import penetration at the regional level. In the next section we turn to how we can use (27) to discipline our empirical analysis and identification strategy.

How does globalization influence U.S. inflation dynamics? According to our model, there are two clear channels through which trade openness affects inflation. Increasing openness, to the extent that it makes cheaper foreign final goods more accessible to U.S. consumers, leads to lower import prices and ultimately consumer prices. This effect leads to increases in import penetration, Λ_{rt}, which in turn lead to lower inflation as reflected on (27). We refer to this effect through changes in import penetration as the *direct effect* of openness on inflation.

In addition to this direct effect, higher import penetration can reduce the sensitivity of inflation to domestic slack—that is, it can lower the slope of the Phillips curve. All else equal, the greater the share of imports in final consumption, the less sensitive consumer prices are to changes in slack. For instance, in an economy that only consumes imported goods, the relation between inflation and domestic slack may be entirely decoupled because consumption is independent of domestic production, and therefore real marginal costs. As previously explained, this mechanism plays out by changing the elasticity of real marginal cost with respect to the unemployment gap. We refer to this channel as the *indirect effect*.

3 Specification and Estimation of the Open-economy Phillips Curve Using Regional Data

We rely on the predictions of the model that we developed in the previous section to discipline our empirical analysis and study the effects of globalization on inflation dynamics. More specifically, our ultimate objective is to estimate an open-economy Phillips curve for the U.S., and then use our estimates to measure the effects of trade openness on U.S. inflation. In this section, we develop our approach to specify and estimate open-economy regional Phillips curves, which can be aggregated to obtain a national-level open-economy Phillips curve.

The estimation of structural Phillips curves in closed-economy settings poses important identification challenges. Many of these arise from endogeneity issues that the literature has clearly identified (Barnichon and Mesters, 2020; Furlanetto and Lepetit, 2025). To circumvent these issues, a relatively new approach in the literature relies on regional data to estimate regional Phillips curves.[18] The estimation and identification strategy that we propose in this paper extends this approach to the case of open-economy structural Phillips curves as specified by equation (27).

The benefits of using regional data for identification extend naturally to the case of open-economy Phillips curves in a monetary union. This result stems from the fact that regional data can help address two prevalent issues in the estimation of both the Phillips curve and the economic consequences of trade: first, the potential lack of variability in aggregate data and, second, the endogeneity issues arising from aggregate confounders in aggregate data. Early work on the labor-

[18] See, for example, McLeay and Tenreyro (2019); Hazell et al. (2022); Fitzgerald et al. (2020).

market effects of trade identified the advantages of using regional data and the literature has widely relied on this approach ever since (Borjas and Ramey, 1995; Chiquiar, 2008; Topalova, 2010; Autor et al., 2013).

To provide a more detailed explanation of how regional data allow us to estimate (27) while addressing endogeneity issues, start by considering the Phillips curve for domestic inflation in (22). Following Hazell et al. (2022), assume that the dynamics of deviations of the unemployment rate from its natural rate can be approximated by means of an AR(1) process and iterate (22) forward to rewrite (27) as

$$\pi_{rt} = -\frac{\kappa_r}{1-\beta\rho_u}\widehat{u}_{rt} + \mathbb{E}_{t+\infty}\pi_{H,t+\infty} - \frac{\Delta\Lambda_{rt}}{\eta - 1}, \tag{28}$$

where $\mathbb{E}_{t+\infty}\pi_{H,t+\infty}$ denotes long-run inflation expectations, which we have assumed are equated across regions in a monetary union—an assumption in line with the literature that estimates regional Phillips curves.[19] Rewriting (27) as (28) allows us to clearly identify how a panel specification with time and region-fixed effects can control for unobservable variables and other aggregate confounders such as monetary policy.

Suppose that we have panel data for inflation, π_{rt}, the unemployment rate, u_{rt}, and import penetration, Λ_{rt}. Consider the following panel specification:

$$\pi_{rt} = \alpha_r + \gamma_t + \delta(\upsilon_r)u_{rt} + \sigma\Delta\Lambda_{rt} + \epsilon_{rt}, \tag{29}$$

where α_r region-fixed effects, γ_t are time-fixed effects, $\Delta\Lambda_{rt}$ denotes the contemporaneous changes in import penetration—as previously defined—$\delta(\cdot)$ is a function of υ_r, and υ_r is the level of region r's import penetration around which we log-linearized our model to obtain equation (27). Relative to the specification typically used for estimation of closed-economy regional Phillips curves, this specification incorporates changes in import penetration, $\Delta\Lambda_{rt}$—capturing the *direct effect* of higher imports reflecting lower import prices and ultimately lower domestic inflation—and a coefficient on unemployment that varies depending on the level of import penetration of each region, $\delta(\upsilon_r)$—capturing the *indirect effect* that a higher import penetration might weaken or strengthen the overall sensitivity of inflation to unemployment.

Comparing (29) to (28), note that time-fixed effects will not only control for aggregate cost-push shocks impacting inflation—which we do not incorporate in our model—but they will also control movements in long-run inflation expectations as pointed out by Hazell et al. (2022).[20] Similarly, time-fixed effects control for endogenous response in a single national monetary policy to shocks, as emphasized by McLeay and Tenreyro (2019) and Fitzgerald et al. (2020). Both, region and

[19] Assuming that $u_{rt+1} - \mathbb{E}_{t+1}[u_{rt+\infty}] = \rho_u\left(u_{rt} - \mathbb{E}_t[u_{rt+\infty}]\right) + \epsilon_{rt}$ implies that we can express $\widehat{u}_{rt+j}$ for any $j \geq 0$ as $\widehat{u}_{rt+j} = \rho_u^j\widehat{u}_{rt} + \sum_{k=1}^{j}\rho_u^{k-1}\epsilon_{rt+k-1}$. Hence, we obtain that domestic price inflation is given by $\pi_{H,rt} = -\kappa_r\mathbb{E}_t\sum_{j=0}^{\infty}\beta^j\widehat{u}_{rt+j} + \mathbb{E}_{t+\infty}\pi_{H,t+\infty}$, where we have assumed that long-run inflation expectations are uniform across regions and $\widehat{u}_{rt} = u_{rt} - \mathbb{E}_t[u_{rt+\infty}]$.

[20] Note that region-fixed effects would capture time-invariant differences in inflation expectations if we had allowed for them.

time-fixed effects will control for time-invariant differences in regional natural output and movements in the aggregate natural rate of unemployment. Lastly, time-fixed effect will also capture changes in import penetration that are common across regions, such as those driven by exchange rate shocks.[21]

Even though a panel specification with fixed effects allows to account for multiple threats to identification, an important endogeneity issue remains given the fact that the relationship between inflation and unemployment specified by (28) is conditional on demand shocks driving changes in unemployment. Addressing this identification issue becomes even more challenging in an open-economy setting because demand-driven changes in unemployment should not be correlated with changes in import penetration. Hence, to properly identify the function $\delta(\cdot)$ and the coefficient σ, we require a set of instruments that capture demand-driven changes in unemployment independent of supply-driven changes in import penetration. For our estimation in Section 4, we will not only consider lagged unemployment to partially control for endogeneity, but we will also rely on a shift-share approach to construct our instruments, as is commonly done in the literature in macroeconomics and international trade that exploits cross-regional variation in data.

To be able to estimate the function $\delta(\upsilon_r)$, we make two additional assumptions. First, even though κ_r is a complicated function of υ_r according to our model, we assume that this function takes the simple form $\delta(\upsilon_r) = \delta + \tau \times \upsilon_r$ with δ and τ parameters to be estimated. Second, we proxy υ_r in any given period t by lagged import penetration, Λ_{rt-1}. The fact that the degree of openness across regions in (27) depends on the time-invariant parameter υ_r stems from the fact that we log-linearized the model around a zero inflation steady state with $P_{F,rt} = P_{rt}$. Hence, in our baseline estimation we allow for more flexibility by permitting the level of import penetration that determines the actual value of δ to vary over time, which is a prominent feature of the data. This assumption allows us to incorporate a time-varying slope of the Phillips curve. We consider a robustness exercise in Section 6 in which we fix regions' openness across time and remain results are qualitatively similar.

Our analysis of how to estimate open-economy regional Phillips curves using regional data has assumed, until now, that we have panel data on inflation, unemployment, and import penetration. We turn now to the data challenges faced to obtain such data for the U.S. representing the monetary union in our model and each of its states representing one of its regions. To create a long panel for U.S. inflation at the state level, we rely on the data recently provided by Hazell et al. (2022), who construct state level CPI prices back to the 1970s using data on micro-level prices by the Bureau of Labor Statistics (BLS). We obtain data on unemployment from the BLS as well. The real data challenge emerges from the fact that the data required to construct a panel going back to the 1970s of import penetration as specified in (26) is not available.[22] Therefore, we propose an methodology to construct a proxy for import penetration in final consumption based on industry-level data. Our methodology is aimed at constructing a measure (i) that is available at the regional level, (ii) that goes back to the 1970s, and (iii) that is consistent across time—that is, that measures the same set

[21] An additional identification issue can arise from assuming that the AR(1) coefficient on unemployment dynamics does not vary across regions. We explore this possibility in a robustness exercise in Section 6.

[22] State-level import data is only available starting in 2008.

of stat characteristics over time. The measure we propose is particularly convenient because it is constructed based on a shift-share design, which addresses certain endogeneity concerns that may arise because of state-level demand shocks correlated with changes in import penetration (Adão et al., 2019).

To construct our measure of import penetration, we use the *Consumption Expenditure Survey* (CEX) to compute final consumption weights across sectors for each state in 1994 (the first year state observations are available), and then use those state-specific weights applied to changes in national import penetration measures at the sectoral level. More specifically, letting j index the sectors that compose the economy and Λ_t^j denote national import penetration for sector j at time t, we construct import penetration for region r as $\Lambda_{rt} \equiv \sum_j \alpha_r^l \Lambda_t^j$. Note that following this procedure implies that all variation across regions in changes in import penetration will come from differences in sectoral final consumption shares. Our approach is similar to other common approaches in international trade that measure regional trade exposure using shift-share designs. We provide further details on how we construct import penetration measures in Section 4.

Based on the insights derived in this section, in the following section we develop the specific empirical methodology that we use to estimate U.S. regional Phillips curves that are consistent with the fact the U.S. is an open economy with varying degrees of openness across states. We then rely on our estimates to quantify the effects of openness on U.S. inflation dynamics and monetary policy trade-offs.

4 Empirical Methodology

4.1 Empirical Specification

Based on the discussion in Section 3, we consider the following baseline specification for our empirical analysis:

$$\pi_{r,t} = \alpha_r + \gamma_t + \delta u_{r,t-1} + \tau u_{r,t-1} IP_{r,t-1} + \sigma \Delta IP_{r,t} + \delta X_{r,t} + \epsilon_{rt}, \tag{30}$$

where $\pi_{r,t}$ is either state-level non-housing CPI inflation (Q4/Q4), or rent inflation; α_r are state fixed effects and γ_t are time fixed effects, $u_{r,t-1}$ is state unemployment in Q4 of previous year; $IP_{r,t-1}$ is lagged state-level import penetration, and $X_{r,t}$ is a set of controls that can vary across state and time and that we incorporate to control for additional confounders. As discussed in Section 3, the τ term in (30) reflects the effects of openness on the slope of regional Phillips curves, while the σ term captures a direct effect of openness on inflation. Moreover, the relationships between structural parameters and the parameters to be estimated in (30) are given by $\kappa_r = -(\delta + \tau IP_{r,t-1})(1 - \beta\rho_u)$ where ρ_u is the AR(1) coefficient on unemployment, and $\eta - 1 = -\frac{1}{\sigma}$.

4.2 Data

We focus on the period post 1976 and gather the data from multiple sources. To construct our measures of inflation, we rely on state-level non-housing CPI inflation (Q4/Q4) from Hazell et al. (2022). From this source we obtain annual data for 31 states from 1977 to 2017. In addition, we collect data on annual rent inflation (only available starting in 1986) for the same states and years from Fair Market Rents from HUD. We collect the data on unemployment from the Quarterly Census of Employment and Wages (QCEW) by the BLS.[23]

The measure of import penetration that is consistent with the model outlined in the previous section should be based on import shares of final consumption in each state. To construct a proxy for this measure, we consider estimates of state-level expenditures at the detailed product level (e.g., cars, men's shirts, jewelry, furniture, etc) for 1994 from the BLS Consumer Expenditure Survey public-use microdata. We also rely on the nominal value of domestic production by NAICS 4-digit industries annually for the period 1976-2017 (for manufacturing sectors, we use the NBER-CES Manufacturing Database, and for agricultural and mining sectors we use BEA detailed GDP-by-Industry data). Lastly, we also collect data on national nominal imports and exports by NAICS 4-digit industry annually from 1976-1989 from Census, 1990 onwards from USITC.

To construct our measures of import penetration at the regional level, we first construct a nationwide measure of import penetration in a given year t for NAICS 4-digit sectors in agriculture, mining, and manufacturing denoted by j as follows:

$$IP_{j,t} = \frac{M_{j,t}}{(Y_{j,t} - X_{j,t}) + M_{j,t}}, \tag{31}$$

where $IP_{j,t}$ is our nationwide measure of import penetration for sector j, $Y_{j,t}$ are national output in sector j, $M_{j,t}$ is national imports in sector j, and $X_{j,t}$ are national imports in sector j, all these measured in period t. Note that this measure can vary from 0 (all consumption is domestic) to 1 (all consumption is foreign-sourced). We then use each state's sectoral expenditure shares from 1994 to form the weighted average across the national sectoral import penetration rates for each year. Let $s_{r,j}$ denote the share of total expenditure on sector j goods by consumers in region r in 1994. Then, our measure of state-specific import penetration is defined as

$$IP_{r,t} = \sum_j s_{r,j} IP_{j,t}. \tag{32}$$

Note that this measure is in line with the measure of import penetration dictated by the model, Λ_{rt}. In particular, $IP_{j,t}$, is also a number between 0 and 1. We compute changes in import penetration as

$$\Delta IP_{r,t} \equiv \sum_j s_{r,j} \Delta IP_{j,t}, \tag{33}$$

[23]The figures in Appendix E show measures of inflation and unemployment rates across U.S. states.

where $\Delta IP_{j,t} = IP_{j,t} - IP_{j,t-1}$.

We now proceed to estimate different versions of equation (30) with the data that we just described.

4.3 OLS Estimates

We start by estimating alternative specifications of equation (30) using non-housing inflation as our measure of inflation. We turn to estimates for rent inflation later in this section and will rely on these estimates to measure the effects of openness on total inflation. Table 1 presents our OLS estimates.

Before estimating the version of specification (30) that is in line with the open-economy Phillips curve that was derived from our model in (27), we consider the case of a closed economy specification and ignore the interaction term between unemployment and lagged import penetration as well as contemporaneous changes in import penetration. Columns (1) and (2) show the results for this case without and with state and time fixed effects, respectively. The estimate in column (1) does not include state and time fixed effects to control for change in inflation expectations. Adding these fixed effects reduces the omitted variable bias in estimating δ in (30) by controlling for changes in these expectations. Our estimate of -0.17 for δ is in line with recent works using regional data and region and time fixed effects to estimate closed-economy regional Phillips curves (Fitzgerald et al., 2020; Hazell et al., 2022).[24]

In line with the theoretical framework that we developed in Section 2, estimates in columns (3) and (4) of Table 1 include the terms that incorporate the effects on inflation of trade openness: the interaction between unemployment and lagged import penetration, $u_{r,t-1}IP_{r,t-1}$, and contemporaneous changes in import penetration, $\Delta IP_{r,t}$. Column (3) shows the estimates for the specification of (30) that aligns perfectly with our model. Turning first to the *indirect effect* of trade openness on inflation—the effects of trade on the slope of the Phillips curve—we estimate the effect of the interaction term between unemployment and lagged import penetration, τ, to be positive and statistically significant. The positive sign of our estimate implies that more open U.S. states tend to have flatter Phillips curves, a relationship that is in line with the predictions of our model under reasonable parametric assumptions, as discussed in Section 2. For our analysis, we have defined import penetration in terms of units of tenths of a percentage point (rather than in terms of 1 percentage point). As an example, consider two regions in a given year, a relatively open region with $IP_r = 0.12$ and a more closed region with $IP_r = 0.08$—a difference of four percentage points. According to our estimates, the unadjusted slope of the Phillips curve, given by $\frac{\partial \pi_{r,t}}{\partial u_{r,t-1}} = \delta + \tau \times (10 \times IP_{r,t-1})$ in the more open region would be -0.11. In the case of the more closed region this slope would be -0.19. The difference between these two slopes is sizable. If we include the lagged level of import penetration as a control, as we do in column (4), we see that our main result regarding the *indirect effect* of trade on inflation dynamics through the changes in the slope of the Phillips curve is unaltered.

[24] See Furlanetto and Lepetit (2025) for a survey of the literature estimating the slope of the Phillips curve.

Turning to the *direct effect* of trade openness on inflation dynamics, our estimates of σ in columns (3) and (4) of Table 1 are negative. This result implies that increases in import penetration in a given year shift the Phillips curve downwards, thus leading to lower inflation. Hence, our estimates are in line with the predictions of our model when the trade elasticity is positive, $\eta - 1 > 0$. However, note that our estimates of the *direct effects of trade* are very imprecise and therefore not statistically significant. There are a number of issues that could be driving the lack of precision in our estimates including potential endogeneity bias. We propose an IV strategy in the following section to address endogeneity issues and provide additional analysis focusing on the *direct effects* in Section 6.

Table 1: Import Penetration and Non-Housing Inflation Rates: OLS Estimates 1986-2017

Dependent variable: Non-housing CPI annual inflation rate in the fourth quarter (Q4/Q4 in % pts)

	(1)	(2)	(3)	(4)
$u_{r,t-1}$	-0.05**	-0.17***	-0.34***	-0.49***
	(0.025)	(0.053)	(0.085)	(0.148)
$u_{r,t-1} IP_{r,t-1}$			0.19**	0.35**
			(0.070)	(0.160)
$\Delta IP_{r,t}$			-1.66	-0.87
			(2.067)	(2.594)
$IP_{r,t-1}$				2.63
				(1.614)
State FE	No	Yes	Yes	Yes
Time FE	No	Yes	Yes	Yes

Notes: $N = 889$. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$; standard errors are clustered by state and time.

4.4 IV Estimates

Even though the fixed effects included in the regressions control for important time-varying aggregate factors—such as monetary policy responses and aggregate inflation expectations—and time-invariant state characteristics, they do not completely rule out the possibility of an endogeneity bias.[25] Ultimately, there still could be omitted factors which contribute to movements in both state-level inflation and state-level unemployment (and also state trade exposure). We propose instrumental variables strategies to deal with these potential issues. In particular, we want to capture regional demand shifts that correlate with unemployment but not with regional supply shocks contained in the error term.

We propose an IV strategy aimed at addressing potential endogeneity bias in both state level unemployment and changes in import penetration.[26] Starting with potential endogeneity bias in the coefficient on state-level unemployment. Our IV strategy builds on the literature on government

[25] See the discussion in Hazell et al. (2022).

[26] Our strategy deviates from that proposed by Hazell et al. (2022) because of the key role played by differential exposure to trade across states.

spending shocks (Nekarda and Ramey (2011), Nakamura and Steinsson (2014), McLeay and Tenreyro (2019)). In particular, we interact state dummy variables with the lagged value of federal defense spending and use these interactions as instruments (which are expected to be informative for state-level unemployment). This is a simpler version of the shift-share strategy used in McLeay and Tenreyro (2019) at the city-level. The identifying assumption is that states are differentially exposed in terms of demand shocks in response to federal defense spending, and that this idiosyncratic state-level variation in demand from federal defense spending is uncorrelated with unobserved state-level supply shocks in the regression residual.

The shift-share procedure we used to construct our measure of import penetration already addresses some of the potential endogeneity of our trade exposure measures (Borusyak et al., 2025). However, we further develop additional instruments based on the strategy in Autor et al. (2013). Specifically, we use both the lagged level and the change in exports from China to the same non-US countries as in Autor et al. (2013) as instruments, which should be informative for the lagged level and change in U.S. import penetration. The identifying assumption of these instruments is that Chinese exports to other countries reflect positive productivity shocks in China and not unobserved U.S. supply shocks.

Table 2: Import Penetration and Non-Housing Inflation Rates: IV Estimates 1986-2017

Dependent variable: Non-housing CPI annual inflation rate in the fourth quarter (Q4/Q4 in % pts)

	(1)	(2)	(3)
$u_{r,t-1}$	-0.19**	-0.56**	-1.03***
	(0.087)	(0.240)	(0.333)
$u_{r,t-1}IP_{r,t-1}$		0.40*	0.87**
		(0.225)	(0.318)
$\Delta IP_{r,t}$		-2.07	0.22
		(8.772)	(8.554)
$IP_{r,t-1}$			5.7**
			(2.667)
State FE	Yes	Yes	Yes
Time FE	Yes	Yes	Yes

Notes: $N = 889$. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$; standard errors are clustered by state and time.

Table 2 presents the estimates that we obtain following our IV strategy. For completeness, column (1) shows our estimate of the slope of the Phillips curve without taking into account the *indirect* and *direct effects* of trade openness. This estimate is comparable to the one we obtained under OLS. Columns (2) and (3) show that our results remain qualitatively similar to our OLS results when we incorporate the effects of trade openness. Focusing on column (2)—our preferred non-housing estimates—we see that the coefficient on the unemployment rate is larger in absolute value, and that the coefficient on the interaction term remains positive and statistically significant. Returning to the example we used with our OLS estimates, our IV results imply that the slope of the Phillips curve of the open region would be -0.08 and that of the more closed would be -0.24.

Moreover, according to our estimates, the typical state in our sample, which opened up to trade during this period and experienced a 10 percentage point increase in import penetration, would have experienced a decline in the slope of its Phillips curve of 0.27.

Turning to our estimates of the *direct effect* of trade openness in column (2) of Table 2, note that while the coefficient on the change in import penetration remains negative, it is still statistically insignificant. There is an important national component in changes in import penetration across states which leads to imprecise estimates. However, taking our IV estimates at face value, we can still compare two state experiencing different changes in import penetration in a given year. For example, for a given t, consider state A with no change in its import penetration, $\Delta IP_{A,t} = 0$, and state B that experiences an increase in its import penetration of one percentage point, $\Delta IP_{B,t} = 0.01$. According to our estimate, all else equal, inflation in state B would be lower than in state A by 0.2 percentage point.[27] Moreover, note equation (27), our estimate $\widehat{\sigma}$ implies an estimate of the trade elasticity, $\widehat{\eta} - 1 \approx 5$, which is consistent with the wide range of estimates for this elasticity in existing literature (Simonovska and Waugh, 2014; Caliendo and Parro, 2015).

4.5 Time-invariant Openness

So far, we have estimated the parameters of interest following specification (30) closely. This specification allows for the level of trade exposure determining the slope of the regional Phillips curves to differ across time periods—that is, $IP_{r,t-1}$ in the interaction term in (30) can vary in each period t—which contrasts with he model's prediction shown in (27) where κ_r is fixed over time. We have followed the previous strategy because of two main reasons. First, theoretically, κ_r in (27) is fixed over time only because we have considered a log-linear approximation of the model around a given steady state. Hence, we believe our empirical strategy is more general because it allows for persistent changes in the level of trade openness. Second, by allowing import exposure to vary over time we exploit additional variation in the data that allows us to better estimate our parameters of interest.

To remain as close as possible to the predictions of our model, we now consider an empirical specification in which the measure of import exposure that we consider in the interaction term of (30) does not change over time. More specifically, we construct the measure of import penetration $\overline{IP}_r$ as the average of import penetration across time in state r, $\overline{IP}_r = \sum_t IP_{r,t}$, and estimate the effects of trade openness on U.S. non-housing inflation by estimating the following specification:

$$\pi_{r,t} = \alpha_r + \gamma_t + \delta u_{r,t-1} + \tau u_{r,t-1}\overline{IP}_r + \sigma \Delta IP_{r,t} + \delta X_{r,t} + \epsilon_{rt}. \tag{34}$$

Table 3 shows the OLS and IV estimates obtained when we consider specification (34). The estimates of δ and τ shown in the first and second rows are in line with our previous result—regions

[27] Note that, in contrast to our estimates of the slope of the Phillips curve, we cannot make statements about the *direct effects* of trade on inflation for an average or typical state at a given point in time because time-fixed effects in our regression also pick up common changes in import penetration across all U.S. states. Under additional assumptions, we revisit this issue in Section 5.3.

that were on average more open over time had flatter Phillips curves. In terms of magnitudes, these estimates are also in line with the one we obtained previously. The median state has an average import penetration of of eight percent. Hence, $\overline{IP}_r = 0.08$ and according to our IV estimates shown in column (2) of Table (34), we obtain that the slope of the non-housing Phillips curve of the median U.S. state was $-0.19 = -1.24 + 1.31 \times (10 \times 0.08)$. Using this measure of trade openness, we obtain the inter-quartile range across states' import exposure is 1.4 percentage point, implying that very open states—those in the highest quartile of the distribution—had basically flat non-housing inflation Phillips curves, while the more closed states—those in the lowest quartile of the distribution—had non-housing inflation Phillips curves with slopes close to -0.38.

Turning to the *direct effect* of trade on U.S. inflation, the fourth row of Table 3 shows that our estimates of σ remain highly imprecise. Interestingly, while the sign of the OLS estimate is in line with our model, the IV estimate we obtain flips signs. That said, both estimates are nor statistically different from zero.

In summary, our result that more open U.S. states should have flatter Phillips curves remains even when we restrict attention to a time-invariant measure of trade exposure interacting with the unemployment rate.

Table 3: Import Penetration and Non-Housing Inflation: IV Estimates with Time-invariant Openness

Dependent variable: Non-housing CPI annual inflation rate in the fourth quarter (Q4/Q4 in % pts)

	OLS Estimates	IV Estimates
	(1)	(2)
$u_{r,t-1}$	-0.80**	-1.24***
	(0.358)	(0.338)
$u_{r,t-1}\overline{IP}_r$	0.79*	1.31***
	(0.433)	(0.487)
$\Delta IP_{r,t}$	-0.55	2.78
	(2.189)	(7.203)

Notes: $N = 889$. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$; standard errors are clustered by state and time. Time period considered: 1986-2017. State- and time-fixed effects included.

4.6 Effects on Rent Inflation

The estimates that we have presented so far are for non-housing inflation. To be able to make statements about the effects of trade openness on total inflation in the U.S., we need to consider rent-inflation which represents approximately 40 percent of CPI inflation. As described in Section 4.2, we complement our data on non-housing inflation across U.S. states with additional data on rental prices to construct rent inflation. We present our estimates of the effects of trade openness on rent inflation in this section.

Both OLS and IV estimates for our specification with rent inflation as the dependent variable are shown in Table 4. Columns (2) and (4) provide OLS and IV estimates, respectively, for our preferred specification. Note that in this case, the *direct* and *indirect* effects of trade on rent inflation are not statistically significant. However, when we consider the specification that excludes the effects of openness on inflation, as we do in columns (1) and (3), we see that there is a negative relationship at the regional level between rent inflation and the unemployment rate. These estimates are in line with housing services being non-tradable goods and, therefore, having prices completely determined by local economic activity. By comparing our estimates for the slope of the Phillips curve in column (3) of Table 4 with those in column (2) of Table 2, we see that the rent-inflation Phillips curve for the average state is steeper than the one for non-housing inflation. This result is in line with those by Hazell et al. (2022).

Table 4: Import Penetration and Rent-Inflation: Estimates 1986-2017

Dependent variable: Rent-inflation rate in the fourth quarter (Q4/Q4 in % pts)

	OLS Estimates		IV Estimates	
	(1)	(2)	(3)	(4)
$u_{r,t-1}$	-0.36*	-0.18	-0.47***	-0.06
	(0.206)	(0.541)	(0.070)	(0.888)
$u_{r,t-1}IP_{r,t-1}$		-0.22		-0.74
		(0.530)		(0.876)
$\Delta IP_{r,t}$		13.9		23.9**
		(9.07)		(13.069)

Notes: $N = 992$. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$; standard errors are clustered by state and time. State- and time-fixed effects included.

4.7 The Slope of the National Phillips Curve

We construct the national Phillips curve slope in each year that is implied by our estimates of the effect of trade.

We can use our estimates of δ and τ to recover an estimate of slope of the national U.S. Phillips curve under certain assumptions. First, we define the time-varying slope of the national non-housing Phillips curve, κ_t, as the average of the slopes of the regional non-housing Phillips curves weighted by each region's population, $\kappa_t \equiv \sum_{r=1}^{I} \nu_{rt}\kappa_{rt}$. In turn, our estimates of these region-specific and time-varying slopes are given by $\kappa_{rt} \equiv -\left(\widehat{\delta} + \widehat{\tau} \times IP_{r,t-1}\right) \times (1 - \beta\rho_u)$, where δ and τ are given by the estimates in column (2) of Table 2. In line with our analysis in Section 3, note that we have adjusted our estimates by $1 - \beta\rho_u$ to reflect the dynamics of the unemployment rate. For this adjustment, we assume $\beta = 0.95$ and estimate $\rho_u = 0.75$ for the unemployment rate from 1986-2017.

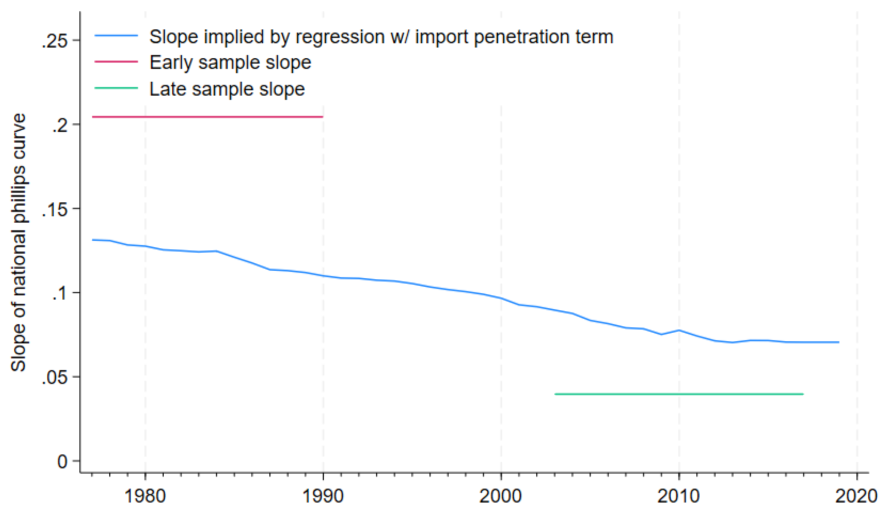
Second, we incorporate our results for rent inflation. Since we do not find clear evidence of any impact of trade openness on the slope of the rent Phillips curve, we use the estimate from the IV regression shown in column (3) of Table 4.[28] To construct the slope of the national Phillips curve,

[28] Note that while maintaining the same estimation period as the rent regression constrained us to an estimation sample starting in the mid 1980s, we can use our estimated coefficients to recover the Phillips curve slope going back to the late 1970s given available trade data.

we simply take a weighted average of the two slopes—of the non-housing and rent Phillips curves—where the weights are 0.6 for non-housing CPI and 0.4 for rent.

The evolution of the slope coefficient of our national Phillips curve can be observed in Figure 3 (blue line). The slope is expressed in terms of the ratio of annual total CPI inflation (Q4/Q4) in percentage points to the prior year's unemployment rate in percentage points. Our baseline results—those that incorporate the effects of trade—imply a national Phillips curve slope of -0.13 in 1977 that declines to about -0.07 in 2017. Moreover, our procedure to construct this slope implies that this flattening is coming entirely from a flattening of the non-housing inflation Phillips curve driven by increased trade openness.

Figure 3: Evolution of the National Phillips Curve Slope Coefficient



Our estimates are comparable to those in existing work estimating these slopes for different periods in time. For instance, Hazell et al. (2022) estimate $\widehat{\delta} = -0.198$ for the pre-1990 period and $\widehat{\delta} = -0.09$ for the post-1990 period.

In summary, increased trade openness by the U.S. has led to a flatter U.S. Phillips curve. Given this result, we turn in the following section to more specific examples of how globalization has affected U.S. inflation dynamics in different episodes and its implications, as well as those of potential de-globalization, on inflation dynamics and monetary policy trade-offs.

5 The Effects of Globalization on U.S. Inflation Dynamics

5.1 How much has trade contributed to a flattening of the U.S. Phillips Curve?

Relying on our estimate of the slope of the national U.S. Phillips curve, we can try to answer the following important question: How much of the flattening of the U.S. Phillips curve can be attributed to increase openness. We will rely on two different approached to provide an answer to this question.

Our first approach relies on comparing the flattening of our estimated open-economy Phillips curve with the flattening of a closed-economy Phillips curve estimated in two different sub-periods of time—an early and a late period. Table 5 provides our OLS and IV estimates for the slope of closed-economy Phillips curves for non-housing and rent inflation for these two sub-period. For non-housing inflation, our early period is 1977-1990 and our later period is 2003-2007. Given that our data for rents only starts in 1986, we consider the sub-period 1986-1999 as the earlier period for rent inflation, and the same late period as with non-housing inflation.

As shown in Figure 3, our IV estimates imply that the slope of the U.S. Phillips curve declined from about -0.20 in 1977-1990 (the red line) to about -0.04 in 2003-2017 (the green line). We interpret these results as showing that the increased exposure of the U.S. economy to imports explains a bit less than 40 percent of the flattening of the U.S. Phillips curve that we estimate occurred over the 1977-2017 period.

Table 5: Phillips Curve Slope: OLS and IV estimates

Dependent variable: CPI annual inflation rate in the fourth quarter (Q4/Q4 in % pts)

	(1)	(2)	(3)	(4)
	A. *Non-housing inflation*			
	1977-1990		2003-2017	
$u_{r,t-1}$	-0.14**	-0.49***	-0.06	-0.01
	(0.055)	(0.150)	(0.086)	(0.071)
State FE	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes
N	231	231	450	450
	B. *Rent inflation*			
	1986-1999		2003-2017	
$u_{r,t-1}$	-0.56	-1.25**	-0.01	-0.20
	(0.348)	(0.464)	(0.328)	(.)
State FE	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes
N	434	434	465	465

Notes: Columns (1) and (3) report OLS estimates, columns (2) and (4) report IV estimates. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$; standard errors are clustered by state and time.

Our second approach is motivated by the fact that our baseline empirical specification (30) only allows for the Phillips curve to flatten over time because of changes in trade openness. However, other forces could have also lead to a flattening of the U.S. Phillips curve over time. We incorporate these other forces into our empirical specification by means of period-specific fixed effects. More specifically, we split the period 1977-2017 into three sub-periods and define sub-period-fixed effects for each of these. The sub-periods we consider are: 1977-1985, 1986-1993, 1994-2005, and 2006-2017. Then, we add an interaction between our sub-period-fixed effects and lagged unemployment to specification (30), and estimate the equation following our IV strategy. Following this approach we can compute the contribution of trade openness to the flattening of the U.S. Phillips Curve under relying on a single empirical specification.

Figure 4: Evolution of the National Phillips Curve Slope Coefficient

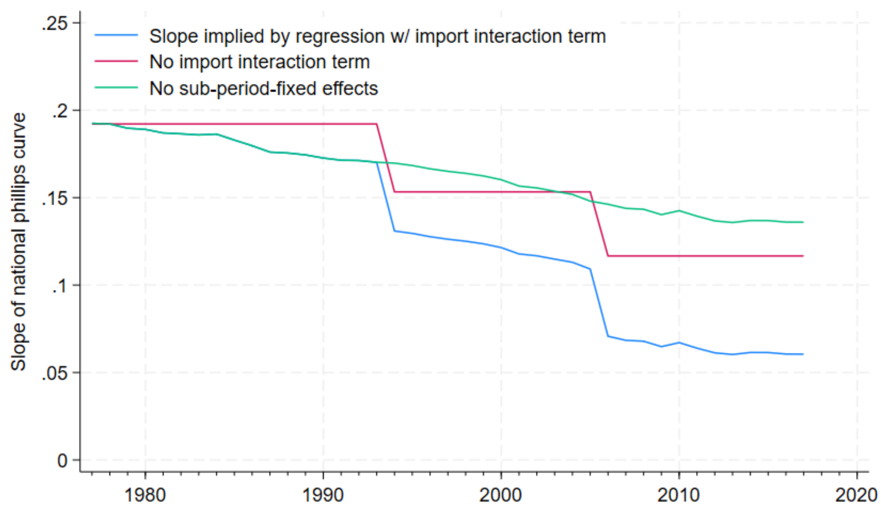


Figure 4 shows that the slope of the Phillips curve is estimated to have declined from 0.19 in 1976 to 0.07 in the past decade and a half (blue lines). If we consider trade openness as the sole factor driving the flattening of the Phillips curve over this period, the green line in Figure 4 shows that increased trade openness made the slope of the Phillips curve decline from 0.19 to 0.14. The latter result suggests that increased openness of the U.S. economy contributed roughly 40 percent of the overall flattening of the U.S. Phillips curve since 1976.

5.2 Counterfactual Analysis "Missing (dis)inflation"

We have established that the U.S. Phillips curve flattened since 1976. A question that remains is if this flattening has quantitatively relevant implications for the macroeconomy. We consider now a simple counterfactual exercise to better understand the implications of the flattening of the U.S. Phillips curve driven by increased trade openness.

We focus on the period 2009-2019 for our counterfactual exercise. A shown in Figure 5a, during this period the U.S. unemployment rate dropped from a maximum level post-2008 of close to 10 percent to a minimum below 4 percent. However, U.S. inflation barely moved during this period, a phenomenon some refer to as "missing disinflation and inflation."[29] We explore how much of the lack of response by inflation to movements in aggregate unemployment, all else constant, can be explained by the flattening of the Phillips curve driven by increased trade.

Given an estimate of the national U.S. unemployment gap over time, we can compute the effect on inflation if the Phillips curve slope had not flattened from exposure to trade. For simplicity, we assume the natural rate of unemployment is constant and equal to the sample average of national

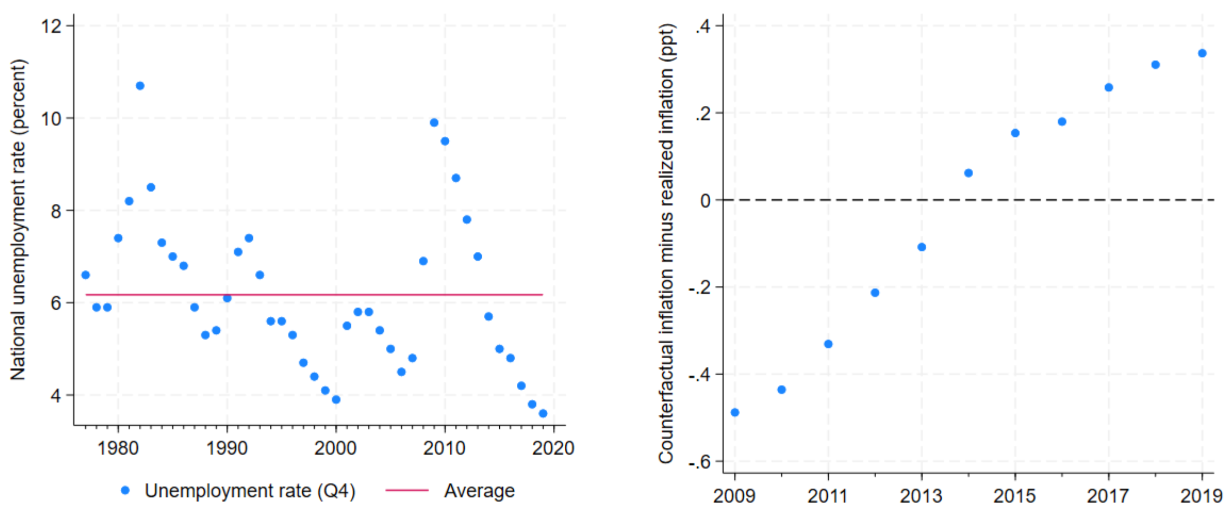
[29] See Forbes (2019) for a detailed account of inflation during this period.

unemployment from 1977 to 2019. Figure 5a plots the data for U.S. unemployment over the period considered as well as its average over the period. Using this estimate for the natural rate of unemployment, we can compute the unemployment gap as the difference between the unemployment rate and this natural rate.

Figure 5: Missing Inflation

(a) National Unemployment

(b) Counterfactual Excess Inflation



With an estimated unemployment gap in hand, we can compute a simple counterfactual where we hold trade fixed at its 1977 level, and thus the Phillips curve slope does not flatten due to trade. The difference between counterfactual inflation and realized inflation in any year, holding all else equal, is just a function of the change in the Phillips curve slope due to trade multiplied by the unemployment gap in that year. The results of this calculation can be seen in Figure 5b.

According to the results of this counterfactual exercise, if the national Phillips curve had not flattened due to trade, then annual headline CPI inflation (Q4/Q4) would have been around 0.5 percentage points lower at the height of the Great Recession, and inflation would have been higher by around 0.3 percentage points in the years just before the global pandemic.

5.3 Average Contribution of Increasing U.S. Openness to Inflation

Our empirical methodology implies that time-fixed effects in our regression also pick up common changes in import penetration across all U.S. states. Therefore, it is not possible to make statements about the direct effects of trade on inflation for an average or typical state at a given point in time unless we make further assumptions about this common component. In this section we provide two estimates of the aggregate direct effect of trade on U.S. inflation relying on two alternative assumptions about this common component.

Our first approach assumes that time-fixed effects do not systematically contain any variation that is common across changes in import penetration across states. In this case, we can provide an estimate of the *direct effect* of trade on non-housing U.S. inflation based on our estimate in column (3) of Table 2. The *direct effect* on total inflation can be computed by taking the change in national import penetration (measured in 0.1 units) in a given year and multiplying by $-1.242(= -2.07 \times 0.6)$. According to the data, national import penetration over the period 1977-2016 grew by 0.23 percentage point annually on average. Hence, our estimates and this average change in national import penetration implies that increased trade directly lowered the annual rate of inflation by around 0.03 percentage points each year. In the years with the highest growth in national import penetration, the annual rate of inflation was reduced by around 0.1 percentage points.

In our baseline empirical methodology, time-fixed effects capture any changes in import penetration that are common across states. Therefore, to correctly estimate the *direct effect* of trade on U.S. inflation, we would ideally tease out the movements in our estimated time-fixed effects related to trade from those that are not. The second approach we consider aims at extracting variation in import penetration that is common across states. We go back to Section 5.1 and consider the specification that incorporates forces other than trade as partial drivers of changes in the slope of Phillips curves over time. Thus, we consider the estimates that we obtain from the specification that includes the interaction between sub-period-fixed effects and the unemployment rate. We obtain an estimate $\widehat{\sigma}$ using our IV strategy and for year-fixed effects, $\widehat{\gamma}_t$. We then estimate by OLS the regression $\widehat{\gamma}_t = \alpha_r + \sigma^{FE}\Delta IP_t + u_t$ for the pre-GFC period 1976-2008—a period over which the U.S. evidently became more integrated with the rest of the world—to obtain an estimate $\widehat{\sigma}^{FE}$. We define the estimated *direct effect* of trade on U.S. inflation as

$$\widehat{\frac{\partial \pi_{US}}{\partial \Delta IP_t}} = (\widehat{\sigma} + \widehat{\sigma}^{FE}) \times 0.6,$$

that is, the effect of a one percentage point change in national import penetration on inflation. Given our estimates $\widehat{\sigma} = -1.49$ and $\widehat{\sigma}^{FE} = -9.06$, we obtain $\widehat{\frac{\partial \pi_{US}}{\partial \Delta IP_t}} = -6.33$, which is 5.1 times larger than our estimate according to the first approach we followed. Hence, from 1976 to 2008–a period of rapid globalization during which national import penetration grew by 0.27 percentage point annually on average–the *direct effect* of trade reduced annual headline CPI inflation around 0.17 percentage point. In the years with the highest growth in import penetration, U.S. annual inflation was reduced by up to 0.4 percentage point.

5.4 Implications of Deglobalization for U.S. Inflation and Monetary Policy Trade-offs

Given our results so far, in this section we explore the implications of deglobalization for U.S. inflation and monetary policy trade-offs. We consider an illustrative deglobalization scenario in which the national-average of our state-level import penetration measure declines from 12.3 percent (its 2017 value) to 6.5 percent (its 1991 value)[30]. This is about a 47 percent decline in import penetration. Such a decline in import penetration could come about as a result of a number of recent trends, such as tariff increases or geopolitical fragmentation.

Through the lens of our analysis, this deglobalization scenario implies both a direct effect on inflation, along with an indirect effect on the slope of the Phillips curve. The size of the direct effect, on a per-year basis, depends on how quickly the import penetration measure declines. If we assume that import penetration takes four years to decline to 6.5 percent, then import penetration declines by around 1.5 percentage points per-year, and given our estimates in the last section, the direct effect of this retrench in globalization is a boost to U.S. headline CPI inflation of about 0.9 ppt. per-year over this period.

Turning to the effects of deglobalization on monetary policy tradeoffs, our analysis suggests that globalization contributed to an increase in the sacrifice ratio from 6.3 in 1976 to 14.5 in the recent period. We rely on the coefficients from the sub-period fixed effects regressions to obtain these numbers. The increase in the sacrifice ratio is, in a sense, a proverbial double-edged sword: When inflation is low, it allows policymakers to support employment with limited risk of overshooting their inflation goal. When inflation is elevated, however, the process of bringing it down entails larger costs in terms of higher unemployment. Turning to our deglobalization scenario described in the previous paragraph, our estimates indicate that the national Phillips curve slope would steepen from about 0.06 to about 0.09. Given this steeper slope, the increase in the unemployment rate gap required to lower inflation 1 percentage point—that is, the sacrifice ratio—would decline from about 16.7 percent previously to about 11.1 percent with lower import penetration.

Thus, there are two conclusions in this analysis: A reversal of globalization would raise inflation along the transition path, but also would reduce the costs of disinflation. The first effect can be sizable, given a large enough and fast enough deglobalization scenario, although transitory. The second effect is permanent and also quantitatively relevant.

6 Robustness Exercises

6.1 Regional Heterogeneity in Unemployment Dynamics

An important assumption that we made when we described our methodology to estimate the parameters in specification (30) was that the persistence of deviations of unemployment from its long-run level did not differ across states. More specifically, in footnote 19, we assumed that ρ_u is the same across states, implying that the persistence of unemployment is the same in all regions. This assumption is convenient because it implies that all heterogeneity in slopes of regional Phillips curves arises from heterogeneity in κ_r. However, this assumption might be at odds with the data. We explore this issue in this section.

[30] We assume that the 2017 value roughly approximates where the U.S. economy has been recently (2024). This seems reasonable since even though our state-specific measures do not extend that recently, cruder national measures such as the goods imports share of GDP at 12.1 percent in 2017 is similar to our measure, and this national share more or less moves sideways after 2017.

According to equation (28), note that the response of inflation to an increase in $\widehat{u}_{rt}$ in region r could be higher because of two different reasons, a steeper Phillips curve or more persistent unemployment dynamics. In the model, the dynamics of unemployment are dictated by demand-side factors that could be correlated with the degree of trade openness. In this case, the heterogeneity that we estimated for regional non-housing Phillips curves could be driven by heterogeneity in unemployment dynamics rather than in trade openness. To investigate this issue, we split states across different levels of average trade exposure, $\overline{IP}_r$, and estimate AR(1) processes for each group of states controlling for state- and quarter-fixed effects. Table 6 shows our estimates, first for the sample pooling all states, and then for states across the four quartiles of the distribution of $\overline{IP}_r$. All point estimates are very similar across groups and, given standard errors, we cannot reject the hypothesis that they are equal. Hence, we conclude that there is no suggestive evidence that unemployment dynamics are systematically correlated with states' level of trade openness.

Table 6: Unemployment Persistence Across Levels of Import Exposure

Dependent variable: Unemployment rate un period t (% pts)

		Level of Trade Exposure			
	(1)	(2)	(3)	(4)	(5)
	All	*Low*	*Mid-low*	*Mid-high*	*High*
$u_{r,t-1}$	0.751	0.756	0.736	0.760	0.714
	(0.032)	(0.040)	(.)	(0.052)	(0.054)
N	992	256	256	256	424

Notes: OLS estimates for the first auto-regressive coefficient of the unemployment rate. All regressions include state- and quarter-fixed effect. Column (1) reports the estimate for the pooled sample, columns (2) through (5) report estimates for sub-samples of states differently exposed to trade. Standard errors are clustered by state and time.

6.2 Alternative Measure of Trade Openness

A key insight of our theoretical framework is that, to study the implications of trade openness for inflation dynamics, we require a measure of trade exposure that reflects the degree of trade openness in final consumption. Such a measure differs from the more common measure of trade exposure used to study the effects of globalization shocks on labor markets (Autor et al., 2013). We now consider this alternative measure of import penetration in our estimation to explore how our results change.

As explained in section 4, our measure of import penetration relies on the shares of expenditure across sectors by final consumers, which we obtained from the BLS Consumer Expenditure Survey public-use microdata for 1994. Our alternative measure—based on the work of Autor et al. (2013)—replaces these shares by the shares of total employment across sectors in each state. More specifically this alternative measure is defined as

$$IP_{r,t}^{M} = \sum_{j} s_{r,j} M_{j,t}, \tag{35}$$

where, as previously stated, $M_{j,t}$ is the level of U.S. imports of sector j goods at time t.[31] As with our baseline measure of import penetration, we compute changes in regions-specific import penetration based on changes in national sector-specific import penetration measures keeping labor shares fixed.

Table 7: Alternative Measure of Import Penetration and Inflation Rates—Phillips Curve: OLS Estimates

Dependent variable: Non-housing CPI annual inflation rate in the fourth quarter (Q4/Q4 in % pts)

	OLS Estimates	IV Estimates
	(1)	(2)
$u_{r,t-1}$	-0.18**	-0.22**
	(0.075)	(0.099)
$u_{r,t-1} IP_{r,t-1}^{M}$	0.40	0.58
	(0.380)	(0.659)
$\Delta IP_{r,t}^{M}$	0.062***	0.419
	(0.015)	(0.393)

Notes: $N = 840$. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$; standard errors are clustered by state and time.

Columns (1) and (2) of Table 7 present OLS and IV estimates using the alternative measure of import penetration, respectively. These estimates can be compared to those shown in column (3) of Table 1 and column (2) of Table 2. Note that independently of the estimation strategy followed, the estimated coefficient for the interaction term has the expected sign, but our estimates are not statistically significant. This result contrasts with the statistically significant coefficients that we obtained using our baseline measure of import penetration. For the case of the effects of changes in import penetration on non-housing inflation, our OLS estimate delivers a positive and statistically significant coefficient. However, when we consider out IV estimate, we see that our estimate is no longer statistically significant. These results highlight the fact that it is key to consider the correct measure of import penetration to understand the effects of trade openness inflation dynamics.

7 Conclusions

We develop a New Keynesian model of an open economy composed of multiple regions that delivers structural equations for regional Phillips curves with slopes that depend on regions' exposure to

[31]To compute the shares, $s_{r,j}$, we used state-level employment in 1990 in tradable NAICS 4-digit industries (about 100 of them) to construct industry employment shares. Then, we apply these 1990 state-level industry employment shares to the national nominal value of imports in these 100 industries for each year-quarter from 1989 Q1 to 2017 Q4.

trade. In line with our model, we construct a new measure of state exposure to international trade in final consumption goods. We exploit variation in this measure across U.S. states to estimate regional Phillips curves.

Our estimates provide evidence that the slopes of state-level Phillips curves for non-housing inflation differ systematically across states. More specifically, we find that more open states exhibit flatter Phillips curves, in line with reasonable parametric assumptions of the model hat we develop. Our point estimates are also consistent with those states that experienced larger contemporaneous increases in import penetration exhibiting lower non-housing inflation.

We further study the effect of trade on U.S. inflation dynamics through a series of counterfactual exercises. We find that the substantial increase in U.S. trade exposure over the past four decades contributed meaningfully to the flattening of both state-level Phillips curves and the aggregate U.S. Phillips curve. Approximately 40 percent of the observed decline in the slope of the national Phillips curve can be attributed to increased import exposure. In addition, we show that absent the trade-induced flattening, inflation dynamics in the U.S. would have been quantitatively different at the height of the Great Recession and during the years immediately preceding the COVID-19 pandemic. Lastly, we show that a sizable reversal of globalization—were it to occur—would constitute a structural shift in inflation dynamics that monetary policy would need to account for.

This flattening of the Phillips curve due to trade is a structural change in the U.S. economy that is quantitatively relevant for the conduct of monetary policy. However, the rapid growth in trade between the United States and the rest of the world that occurred over the previous four decades stalled around the late 2000s. Indeed, it's possible that changes in government policies or firm behavior in the coming years could lead to declining openness of the U.S. economy. Our results suggest that a large change in the openness of the U.S. economy would have important effects on the dynamics of U.S. inflation.

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Appendix A Problem of a Firm Resetting Its Price

Nominal profits by firm i in period t are given by

$$\begin{aligned}\Pi_{rt}(i) &= P_{H,rt}(i)\,Y_{rt}(i) - W_{rt}N_{rt}(i)\\ &= P_{H,rt}(i)\,Y_{rt}(i) - MC_{rt}P_{H,rt}Y_{rt}(i)\end{aligned}$$

where the second equality follows from our definition of real marginal cost, $MC_{rt} \equiv \frac{W_{rt}}{P_{H,rt}A_{rt}}$. Hence, a firm resetting its price in period t will solve

$$\max_{P_{H,rt}(i)} \sum_{k=0}^{\infty} \theta^k \mathbb{E}_t \left[Q_{t,t+k}\Pi_{rt+k}(i)\right]$$

subject to

$$Y_{rt}(i) = \left(\frac{P_{H,rt}(i)}{P_{H,rt}}\right)^{-\varepsilon} Y_{H,rt}.$$

Substituting the last constraint into $\Pi_{rt+k}(i)$ for any $k > 0$ and differentiating with respect to $P_{H,rt}(i)$ delivers:

$$\begin{aligned}\frac{\partial \Pi_{rt+k}(i)}{\partial P_{H,rt}(i)} &= P_{H,rt+k}Y_{H,rt+k}\frac{\partial}{\partial P_{H,rt}(i)}\left[\frac{P_{H,rt}(i)}{P_{H,rt+k}}\left(\frac{P_{H,rt}(i)}{P_{H,rt+k}}\right)^{-\varepsilon} - MC_{rt+k}\left(\frac{P_{H,rt}(i)}{P_{H,rt+k}}\right)^{-\varepsilon}\right]\\ &= Y_{H,rt+k}\left[(1-\varepsilon)\left(\frac{P_{H,rt}(i)}{P_{H,rt+k}}\right)^{-\varepsilon} + \varepsilon MC_{rt+k}\frac{P_{H,rt+k}}{P_{H,rt}(i)}\left(\frac{P_{H,rt}(i)}{P_{H,rt+k}}\right)^{-\varepsilon}\right]\\ &= P_{H,rt}(i)^{-\varepsilon-1}\,Y_{H,rt+k}\left[(1-\varepsilon)\,P^{\varepsilon}_{H,rt+k}P_{H,rt}(i) + \varepsilon MC_{rt+k}P^{1+\varepsilon}_{H,rt+k}\right].\end{aligned}$$

Hence, we have that the first order conditions of a firm resetting prices are given by:

$$\begin{aligned}&\sum_{k=0}^{\infty}\theta^k\mathbb{E}_t\left[Q_{t,t+k}\frac{\partial \Pi_{rt+k}(i)}{\partial P_{H,rt}(i)}\right]\\ =&\sum_{k=0}^{\infty}\theta^k\mathbb{E}_t\left[Q_{t,t+k}P_{H,rt}(i)^{-\varepsilon-1}\,Y_{H,rt+k}\left\{(1-\varepsilon)\,P^{\varepsilon}_{H,rt+k}P_{H,rt}(i) + \varepsilon MC_{rt+k}P^{1+\varepsilon}_{H,rt+k}\right\}\right] = 0,\end{aligned}$$

which implies that the optimal price must be such that

$$\sum_{k=0}^{\infty}(\theta\beta)^k\mathbb{E}_t\left[\frac{U'(C_{rt+k})}{P_{rt+k}}Y_{H,rt+k}\left\{(1-\varepsilon)\,P^{\varepsilon}_{H,rt+k}P_{H,rt}(i) + \varepsilon MC_{rt+k}P^{1+\varepsilon}_{H,rt+k}\right\}\right] = 0 \qquad \text{(A.36)}$$

where we have used the fact that $Q_{t,t+k} = \beta^k \frac{U'(C_{rt+k})}{U'(C_{rt})}\frac{P_{rt}}{P_{rt+k}}$.

Solving (A.36) for $P_{H,rt}(i)$ delivers the expression for the optimal price in (19). Note that all firms resetting prices at t in region r will set the same price. Call this price $P^R_{H,rt}$, and note that $P^R_{H,rt} = \frac{\varepsilon}{\varepsilon-1}\frac{G^1_{rt}}{G^2_{rt}}$ where

$$\begin{aligned} G^1_{rt} &= \sum_{k=0}^{\infty} (\theta\beta)^k \, \mathbb{E}_t \left\{ \frac{U'(C_{rt+k})}{P_{rt+k}} MC_{rt+k} P^{1+\varepsilon}_{H,rt+k} Y_{H,rt+k} \right\} \\ &= \frac{U'(C_{rt})}{P_{rt}} MC_{rt} P^{1+\varepsilon}_{H,rt} Y_{H,rt+k} + \theta\beta \mathbb{E}_t \left\{ G^1_{rt+1} \right\} \end{aligned}$$

and

$$\begin{aligned} G^2_{rt} &= \sum_{k=0}^{\infty} (\theta\beta)^k \, \mathbb{E}_t \left\{ \frac{U'(C_{rt+k})}{P_{rt+k}} P^{\varepsilon}_{H,rt+k} Y_{H,rt+k} \right\} \\ &= \frac{U'(C_{rt})}{P_{rt}} P^{\varepsilon}_{H,rt} Y_{H,rt+k} + \theta\beta \mathbb{E}_t \left\{ G^2_{rt+1} \right\}. \end{aligned}$$

Define the following auxiliary variables, $g^1_{rt} \equiv \frac{G^1_{rt}}{P^{1+\varepsilon}_{H,rt}}$ and $g^2_{rt} \equiv \frac{G^2_{rt}}{P^{\varepsilon}_{H,rt}}$, and note that

$$\begin{aligned} g^1_{rt} &= \frac{U'(C_{rt})}{P_{rt}} MC_{rt} Y_{H,rt+k} + \theta\beta \mathbb{E}_t \left\{ g^1_{rt+1} \left(1+\pi_{H,rt+1}\right)^{1+\varepsilon} \right\}, \\ g^2_{rt} &= \frac{U'(C_{rt})}{P_{rt}} Y_{H,rt+k} + \theta\beta \mathbb{E}_t \left\{ g^2_{rt+1} \left(1+\pi_{H,rt+1}\right)^{\varepsilon} \right\}, \\ \frac{G^1_{rt}}{G^2_{rt}} &= P_{H,rt} \frac{g^1_{rt}}{g^2_{rt}}. \end{aligned}$$

Hence, we can write the reset price as

$$P^R_{H,rt} = \frac{\varepsilon}{\varepsilon - 1} P_{H,rt} \frac{g^1_{rt}}{g^2_{rt}},$$

which implies that

$$\left(1 + \pi^R_{H,rt}\right) = \frac{\varepsilon}{\varepsilon - 1} \left(1 + \pi_{H,rt}\right) \frac{g^1_{rt}}{g^2_{rt}} \tag{A.37}$$

where $\pi^R_{H,rt} \equiv \frac{P^R_{H,rt}}{P_{H,rt-1}} - 1$ and $\pi_{H,rt} \equiv \frac{P_{H,rt}}{P_{H,rt-1}}$. Moreover, because of our Calvo assumption, we have that reset prices and domestic prices are related by the following equation,

$$P^{1-\varepsilon}_{H,rt} = (1-\theta) P^R_{H,rt} + \theta P^{1-\varepsilon}_{H,rt-1},$$

which implies that

$$\left(1 + \pi_{H,rt}\right)^{1-\varepsilon} = (1-\theta)\left(1 + \pi^R_{H,rt}\right)^{1-\varepsilon} + \theta. \tag{A.38}$$

Log-linearizing equations (A.37 and (A.38), we obtain

$$\widetilde{\pi}^R_{H,rt} = \widetilde{\pi}_{H,rt} + \widetilde{g}^1_{rt} - \widetilde{g}^2_{rt},$$
$$\widetilde{\pi}^R_{H,rt} = \frac{\widetilde{\pi}_{H,rt}}{1-\theta}.$$

implying that

$$\widetilde{\pi}_{H,rt} = \frac{1-\theta}{\theta}\left(\widetilde{g}^1_{rt} - \widetilde{g}^2_{rt}\right)$$

where $\widetilde{\pi}^R_{H,rt} \equiv \pi^R_{H,rt} - \pi^R_{H,r}$, $\widetilde{\pi}_{H,rt} \equiv \pi_{H,rt} - \pi_{H,r}$, and $\tilde{g}^i_{rt} \equiv \log g^i_{rt} - \log g^i_r$ for $i = 1, 2$ denote deviations from steady state variables.

Log-linearizing the auxiliary variables around a zero inflation steady state, we obtain that

$$\widetilde{g}^1_{rt} = (1-\sigma)(1-\theta\beta)\tilde{c}_{rt} + (\eta-1)(1-\theta\beta)\tilde{p}_{rt} + (1-\theta\beta)\widetilde{mc}_{rt} + (1+\varepsilon-\eta)\theta\beta\mathbb{E}_t\widetilde{\pi}_{H,rt+1} + \theta\beta\mathbb{E}_t\widetilde{g}^1_{rt+1},$$

and similarly,

$$\widetilde{g}^2_{rt} = (1-\sigma)(1-\theta\beta)\tilde{c}_{rt} + (\eta-1)(1-\theta\beta)\tilde{p}_{rt} + (\varepsilon-\eta)\theta\beta\mathbb{E}_t\widetilde{\pi}_{H,rt+1} + \theta\beta\mathbb{E}_t\widetilde{g}^2_{rt+1},$$

where variables defined by small letters refer to the log of variables defined in terms of capital letters, and tilde denotes log deviations from a zero-inflation steady state. Therefore,

$$\begin{aligned}
\frac{\theta}{1-\theta}\widetilde{\pi}_{H,rt} &= \widetilde{g}^1_{rt} - \widetilde{g}^2_{rt+1} \\
&= (1-\theta\beta)\widetilde{mc}_{rt} + \theta\beta\mathbb{E}_t\widetilde{\pi}_{H,rt+1} + \theta\beta\mathbb{E}_t\left(\widetilde{g}^1_{rt+1} - \widetilde{g}^2_{rt+1}\right) \\
&= (1-\theta\beta)\widetilde{mc}_{rt} + \theta\beta\mathbb{E}_t\widetilde{\pi}_{H,rt+1} + \theta\beta\mathbb{E}_t\left(\widetilde{\pi}^R_{H,rt+1} - \widetilde{\pi}_{H,rt+1}\right) \\
&= (1-\theta\beta)\widetilde{mc}_{rt} + \theta\beta\mathbb{E}_t\widetilde{\pi}_{H,rt+1} + \theta\beta\mathbb{E}_t\left(\frac{\theta}{1-\theta}\widetilde{\pi}_{H,rt+1}\right)
\end{aligned}$$

which implies that

$$\widetilde{\pi}_{H,rt} = \frac{(1-\theta\beta)(1-\theta)}{\theta}\widetilde{mc}_{rt} + \beta\mathbb{E}_t\left[\widetilde{\pi}_{H,rt+1}\right],$$

which delivers equation (20) around a zero-inflation steady state.

Appendix B Proof of Lemma 1

Consider our definition of real marginal cost, $MC_{rt} \equiv \frac{W_{rt}}{A_{rt}P_{H,rt}}$, and note that

$$\begin{aligned} MC_{rt} &= \frac{W_{rt}}{A_{rt}P_{rt}}\frac{P_{rt}}{P_{H,rt}} \\ &= \frac{1}{A_{rt}}C_{rt}^{\sigma}N_{rt}^{\varphi}\frac{P_{rt}}{P_{H,rt}} \\ &= \frac{1}{A_{rt}}C_{rt}^{\sigma}N_{rt}^{\varphi}\left((1-\upsilon_r)+\upsilon_r S_{rt}^{1-\eta}\right)^{\frac{1}{1-\eta}}, \end{aligned}$$

where the second equality follows from the intratemporal condition (13) and the third equality follows after rewriting $\frac{P_{rt}}{P_{H,rt}}$ in terms of the terms of trade, S_{rt}.

We start by rewriting C_{rt} and N_{rt} in terms of region r's output and foreign output. From the intertemporal optimality conditions (14) and the assumption of a complete international asset markets, we have that

$$\beta\left(\frac{C_{t+1}^*}{C_t^*}\right)^{-\sigma}\left(\frac{P_t^*}{P_{t+1}^*}\right)\left(\frac{\mathcal{E}_t}{\mathcal{E}_{t+1}}\right)=Q_{t,t+1},$$

and

$$\beta\left(\frac{C_{rt+1}}{C_{rt}}\right)^{-\sigma}\left(\frac{P_{rt}}{P_{rt+1}}\right)=Q_{t,t+1},$$

for every region r, implying that, under suitable normalizations, the following condition must hold for all regions r:

$$C_{rt}=C_t^*\left(Z_{rt}\right)^{\frac{1}{\sigma}}\left(Q_{rt}\right)^{\frac{1}{\sigma}} \tag{B.39}$$

for all t, where $Q_{rt} \equiv \frac{\mathcal{E}_t P_t^*}{P_{rt}} = \frac{P_{F,t}}{P_{rt}}$ denotes the real exchange rate for region r. Moreover, given our assumption of each region r being a small open economy, we have that $C_t^* = Y_t^*$ where Y_t^* denotes foreign output. Hence,

$$C_{rt}=Y_t^*\left(Z_{rt}\right)^{\frac{1}{\sigma}}\left(Q_{rt}\right)^{\frac{1}{\sigma}}.$$

Turning to N_{rt}, note that the labor market clearing condition, $N_{rt}=\int_0^1 N_{rt}(i)\,di$, and market clearing of intermediate goods produced in r imply that

$$A_{rt}\int_0^1 N_{rt}(i)\,di = A_{rt}N_{rt} = \int_0^1\left(C_{H,rt}(i)+X_{rt}(i)\right)di = Y_{H,rt}\int_0^1\left(\frac{P_{H,rt}(i)}{P_{H,rt}}\right)^{-\varepsilon}di,$$

where $Y_{H,rt} \equiv C_{H,rt} + X_{rt}$. Defining the following measure of price dispersion,

$$\varrho_{H,rt} = \int_0^1 \left(\frac{P_{H,rt}(i)}{P_{H,rt}} \right)^{-\varepsilon} di, \tag{B.40}$$

we obtain the aggregate production function for goods produced domestically in region r:

$$Y_{H,rt} = \frac{A_{rt} N_{rt}}{\varrho_{H,rt}}.$$

Hence, we can rewrite real marginal cost as

$$\begin{aligned} MC_{rt} &= \frac{1}{A_{rt}} C_{rt}^{\sigma} N_{rt}^{\varphi} \left((1 - \upsilon_r) + \upsilon_r S_{rt}^{1-\eta} \right)^{\frac{1}{1-\eta}} \\ &= \frac{1}{A_{rt}} \left(Y_t^* \right)^{\sigma} Z_{rt} Q_{rt} \left(\frac{\varrho_{H,rt} Y_{H,rt}}{A_{rt}} \right)^{\varphi} \left((1 - \upsilon_r) + \upsilon_r S_{rt}^{1-\eta} \right)^{\frac{1}{1-\eta}} \\ &= A_{rt}^{-(1+\varphi)} \left(Y_t^* \right)^{\sigma} Z_{rt} S_{rt} \left(\varrho_{H,rt} Y_{H,rt} \right)^{\varphi}, \end{aligned} \tag{B.41}$$

where the third equality follows from the fact that the real exchange rate, $Q_{rt} \equiv \frac{P_{F,rt}}{P_{rt}}$, can be expressed as

$$Q_{rt} = S_{rt} \frac{P_{H,rt}}{P_{rt}} = S_{rt} \left((1 - \upsilon_r) + \upsilon_r S_{rt}^{1-\eta} \right)^{\frac{-1}{1-\eta}}.$$

Consider now the market clearing condition for tradable intermediates. We have that:

$$\begin{aligned} Y_{H,rt} &= C_{H,rt} + X_{rt} \\ &= (1 - \upsilon_r) \left(\frac{P_{H,rt}}{P_{rt}} \right)^{-\eta} C_{rt} + \frac{\upsilon^*}{I} S_{rt}^{\eta} Y_t^* \\ &= \left(\frac{P_{H,rt}}{P_{rt}} \right)^{-\eta} C_{rt} \left[(1 - \upsilon_r) + \frac{\upsilon^*}{I} \left(\frac{P_{H,rt}}{P_{rt}} \right)^{\eta} S_{rt}^{\eta} \frac{Y_t^*}{C_{rt}} \right] \\ &= \left(\frac{P_{H,rt}}{P_{rt}} \right)^{-\eta} C_{rt} \left[(1 - \upsilon_r) + \frac{\upsilon^*}{I} \left(Z_{rt} \right)^{-\frac{1}{\sigma}} Q_{rt}^{\eta - \frac{1}{\sigma}} \right] \\ &= \left((1 - \upsilon_r) + \upsilon_r S_{rt}^{1-\eta} \right)^{\frac{\eta}{1-\eta}} C_{rt} \left[(1 - \upsilon_r) + \frac{\upsilon^*}{I} \left(Z_{rt} \right)^{-\frac{1}{\sigma}} Q_{rt}^{\eta - \frac{1}{\sigma}} \right], \end{aligned}$$

where the second equality follows from (11) and (6), the fourth equality follows from (B.39), and the fifth equality rewrites relative prices in terms of terms of trade relying on equation (12). Define

$$\phi_r \equiv \frac{(1 - \upsilon_r)}{(1 - \upsilon_r) + \frac{\upsilon^*}{I}}. \tag{B.42}$$

By taking logs and relying on first order approximations, the previous market clearing condition can be expressed as

$$
\begin{aligned}
y_{H,rt} &= c_{rt} + \eta \upsilon_r s_{rt} - (1-\phi_r)\frac{1}{\sigma} z_{rt} + (1-\phi_r)\left(\eta - \frac{1}{\sigma}\right) q_{rt} \\
&= c_{rt} + \eta \upsilon_r s_{rt} - (1-\phi_r)\frac{1}{\sigma} z_{rt} + (1-\phi_r)\left(\eta - \frac{1}{\sigma}\right)(1-\upsilon_r) s_{rt} \\
&= c_{rt} - (1-\phi_r)\frac{1}{\sigma} z_{rt} + \frac{\upsilon_r \left(\sigma\eta + (1-\upsilon_r)(\sigma\eta - 1)\left(\frac{1-\phi_r}{\upsilon_r}\right)\right)}{\sigma} s_{rt} \\
&= c_{rt} - (1-\phi_r)\frac{1}{\sigma} z_{rt} + \frac{\upsilon_r \omega_r}{\sigma} s_{rt}
\end{aligned}
$$

where

$$
\omega_r \equiv \sigma\eta + (1-\upsilon_r)(\sigma\eta - 1)\left(\frac{1-\phi_r}{\upsilon_r}\right), \tag{B.43}
$$

and where the second equality relies on our definition of real exchange rate, which up to a first order approximation is related to the terms of trade by

$$
q_{rt} = (1-\upsilon_r) s_{rt}. \tag{B.44}
$$

Combining the previous equation with condition (B.39), which in logs is given by

$$
c_{rt} = y_t^* + \frac{1}{\sigma} z_{rt} + \frac{1}{\sigma} q_{rt},
$$

and with equation (B.44), we obtain that

$$
\begin{aligned}
y_{H,rt} &= c_{rt} - (1-\phi_r)\frac{1}{\sigma} z_{rt} + \frac{\upsilon_r \omega_r}{\sigma} s_{rt} \\
&= y_t^* + \frac{1}{\sigma} z_{rt} + \frac{1}{\sigma} q_{rt} - (1-\phi_r)\frac{1}{\sigma} z_{rt} + \frac{\upsilon_r \omega_r}{\sigma} s_{rt} \\
&= y_t^* + \frac{\phi_r}{\sigma} z_{rt} + \frac{(1-\upsilon_r) + \upsilon_r \omega_r}{\sigma} s_{rt}.
\end{aligned}
$$

Hence, we can express the terms of trade in terms of outputs and demand shocks:

$$
s_{rt} = \sigma_r^{\upsilon}(y_{rt} - y_t^*) - \sigma_r^{\upsilon}\frac{\phi_r}{\sigma} z_{rt}, \tag{B.45}
$$

where we have defined $y_{rt} \equiv y_{H,rt}$ to simplify notation, and

$$
\sigma_r^{\upsilon} \equiv \frac{\sigma}{(1-\upsilon_r) + \upsilon_r \omega_r}. \tag{B.46}
$$

Going back to the expression for real marginal cost, equation (B.41), in terms of logs where

$mc_{rt} = \log MC_{rt}$, this equation is given by

$$mc_{rt} = \sigma y_t^* + \varphi y_{rt} + z_{rt} - (1+\varphi) a_{rt} + s_{rt}, \tag{B.47}$$

where we have used the fact that $\log [\varrho_{H,rt}] \approx 0$ for a first-order log-linear approximation. Substituting (B.45) into this last expression we get that

$$\begin{aligned} mc_{rt} &= \sigma y_t^* + \varphi y_{rt} + z_{rt} - (1+\varphi) a_{rt} + s_{rt} \\ &= \sigma y_t^* + \varphi y_{rt} + z_{rt} - (1+\varphi) a_{rt} + \sigma_r^{\upsilon} (y_{rt} - y_t^*) - \sigma_r^{\upsilon} \frac{\phi_r}{\sigma} z_{rt} \\ &= (\sigma - \sigma_r^{\upsilon}) y_t^* + (\varphi + \sigma_r^{\upsilon}) y_{rt} - (1+\varphi) a_{rt} + \left(1 - \sigma_r^{\upsilon} \frac{\phi_r}{\sigma}\right) z_{rt}. \end{aligned}$$

Let $x_{rt} = y_{rt} - y_{rt}^n$ denote output deviations from its natural level. Note that y_{rt}^n must be such that $mc_{rt} = -\log \frac{\varepsilon}{\varepsilon-1}$. Let $\mu \equiv \log \frac{\varepsilon}{\varepsilon-1}$. Then, the natural level of output must be such that

$$(\varphi + \sigma_r^{\upsilon}) y_{rt}^n = -\mu - (\sigma - \sigma_r^{\upsilon}) y_t^* - \left(1 - \sigma_r^{\upsilon} \frac{\phi_r}{\sigma}\right) z_{rt} + (1+\varphi) a_{rt},$$

implying that

$$\begin{aligned} y_{rt}^n &= \frac{-\mu}{\varphi + \sigma_r^{\upsilon}} - \left(\frac{\sigma - \sigma_r^{\upsilon}}{\varphi + \sigma_r^{\upsilon}}\right) y_t^* - \frac{1 - \sigma_r^{\upsilon} \frac{\phi_r}{\sigma}}{\varphi + \sigma_r^{\upsilon}} z_{rt} + \frac{1+\varphi}{\varphi + \sigma_{\upsilon}} a_{rt} \\ &= \frac{-\mu}{\varphi + \sigma_r^{\upsilon}} + \frac{1+\varphi}{\varphi + \sigma_r^{\upsilon}} a_{rt} + \frac{\sigma_r^{\upsilon} - \sigma}{\varphi + \sigma_r^{\upsilon}} y_t^* - \frac{1 - \sigma_r^{\upsilon} \frac{\phi_r}{\sigma}}{\varphi + \sigma_r^{\upsilon}} z_{rt}, \end{aligned}$$

which delivers the expression for the natural level of output. Moreover, note that

$$\begin{aligned} mc_{rt} - mc_r &= \left((\sigma - \sigma_r^{\upsilon}) y_t^* + (\varphi + \sigma_r^{\upsilon}) y_{rt} - (1+\varphi) a_{rt} + \left(1 - \sigma_r^{\upsilon} \frac{\phi_r}{\sigma}\right) z_{rt}\right) \\ &\quad - \left((\sigma - \sigma_r^{\upsilon}) y_t^* + (\varphi + \sigma_r^{\upsilon}) y_{rt}^n - (1+\varphi) a_{rt} + \left(1 - \sigma_r^{\upsilon} \frac{\phi_r}{\sigma}\right) z_{rt}\right) \\ &= (\varphi + \sigma_r^{\upsilon}) x_{rt}, \end{aligned}$$

and that

$$\begin{aligned} x_{rt} &= y_{rt} - y_{rt}^n \\ &= a_{rt} + n_{rt} - \log (\varrho_{H,rt}) - (a_{rt} + n_{rt}^n) \\ &= n_{rt} - \log (\varrho_{H,rt}) - n_{rt}^n \\ &= -(u_{rt} - u_{rt}^n), \end{aligned}$$

where we use again the fact that we are considering a first-order approximation around a zero-inflation steady state.

Consider the slope of region r's Phillips curve given by (B.46). Note that only the denominator of this expression depends on regions' level of openness to the rest of the world, υ_r. We analyze how this denominator changes as openness, υ_r, varies. Let us consider the derivative of the denominator of (B.46),

$$\frac{d}{d\upsilon_r}\left[(1-\upsilon_r)+\upsilon_r\omega_r\right]=-1+\left(\omega_r+\upsilon_r\frac{d\omega_r}{d\upsilon_r}\right). \tag{B.48}$$

If this derivative has a positive sign, then a region that becomes more open would experience a decline in the slope of its Phillips curve.

Starting with the derivative of the term ω_r, we have that

$$\begin{aligned}
\upsilon_r\frac{d\omega_r}{d\upsilon_r}&=\upsilon_r\frac{d}{d\upsilon_r}\left[\sigma\eta+(\sigma\eta-1)(1-\upsilon_r)\left(\frac{1-\phi_r}{\upsilon_r}\right)\right]\\
&=\upsilon_r(\sigma\eta-1)\left[-\frac{d\phi_r}{d\upsilon_r}\left(\frac{1-\phi_r}{\upsilon_r}\right)+(1-\upsilon_r)\left(\frac{-\frac{d\phi_r}{d\upsilon_r}\upsilon_r-(1-\phi_r)}{\upsilon_r^2}\right)\right]\\
&=(\sigma\eta-1)\left[-\frac{d\phi_r}{d\upsilon_r}(1-\phi_r)+(1-\upsilon_r)\left(-\frac{d\phi_r}{d\upsilon_r}-\frac{(1-\phi_r)}{\upsilon_r}\right)\right]\\
&=-(\sigma\eta-1)\left[\frac{d\phi_r}{d\upsilon_r}(2-\phi_r-\upsilon_r)+\frac{(1-\upsilon_r)(1-\phi_r)}{\upsilon_r}\right]
\end{aligned} \tag{B.49}$$

Given the definition of ϕ_r in (B.42), we compute its derivative as follows. Let $D=(1-\upsilon_r)+\frac{\upsilon^*}{I}$. Then,

$$\begin{aligned}
\frac{d\phi_r}{d\upsilon_r}&=\frac{d}{d\upsilon_r}\left[\frac{D-\frac{\upsilon^*}{I}}{D}\right]=\frac{-D+(1-\upsilon_r)}{D^2}\\
&=\frac{-\left((1-\upsilon_r)+\frac{\upsilon^*}{I}\right)+(1-\upsilon_r)}{\left((1-\upsilon_r)+\frac{\upsilon^*}{I}\right)^2}\\
&=\frac{\phi_r}{(1-\upsilon_r)}\frac{-\frac{\upsilon^*}{I}}{(1-\upsilon_r)+\frac{\upsilon^*}{I}}\\
&=-\frac{\phi_r(1-\phi_r)}{(1-\upsilon_r)}
\end{aligned}$$

Substituting the last expression into (B.49), we have that

$$\begin{aligned}
\upsilon_r\frac{d\omega_r}{d\upsilon_r}&=-(\sigma\eta-1)\left[\frac{d\phi_r}{d\upsilon_r}(2-\phi_r-\upsilon_r)+\frac{(1-\upsilon_r)(1-\phi_r)}{\upsilon_r}\right]\\
&=(\sigma\eta-1)\left[\frac{\phi_r(1-\phi_r)}{(1-\upsilon_r)}(2-\phi_r-\upsilon_r)-\frac{(1-\upsilon_r)(1-\phi_r)}{\upsilon_r}\right]\\
&=(1-\upsilon_r)(\sigma\eta-1)\left(\frac{1-\phi_r}{\upsilon_r}\right)\left[\frac{\phi_r(1-\phi_r)\upsilon_r}{(1-\upsilon_r)^2}+\frac{\phi_r(1-\upsilon_r)\upsilon_r}{(1-\upsilon_r)^2}-1\right].
\end{aligned}$$

We go back now to (B.48) and obtian that

$$
\begin{aligned}
\frac{d}{d\upsilon_r}\left[(1-\upsilon_r)+\upsilon_r\omega_r\right] &= -1+\left(\omega_r+\upsilon_r\frac{d\omega_r}{d\upsilon_r}\right)\\
&= (\sigma\eta-1)+(1-\upsilon_r)(\sigma\eta-1)\left(\frac{1-\phi_r}{\upsilon_r}\right)\\
&+(1-\upsilon_r)(\sigma\eta-1)\left(\frac{1-\phi_r}{\upsilon_r}\right)\left[\frac{\phi_r(1-\phi_r)\upsilon_r}{(1-\upsilon_r)^2}+\frac{\phi_r(1-\upsilon_r)\upsilon_r}{(1-\upsilon_r)^2}-1\right]\\
&= (\sigma\eta-1)\left(1+\frac{(1-\upsilon_r)(1-\phi_r)}{\upsilon_r}\left(\frac{\phi_r(1-\phi_r)\upsilon_r}{(1-\upsilon_r)^2}+\frac{\phi_r(1-\upsilon_r)\upsilon_r}{(1-\upsilon_r)^2}\right)\right)\\
&= (\sigma\eta-1)\left(1+(1-\phi_r)\left(\frac{\phi_r(1-\phi_r)}{(1-\upsilon_r)}+\phi_r\right)\right).
\end{aligned}
$$

Given that $\upsilon_r < 1$ and $\phi_r < 1$, this expression is positive as long as $\sigma\eta > 1$. Hence, more open regions will have flatter Phillips curves as long as $\sigma\eta > 1$.

Appendix C A Model with Intermediate Inputs

Suppose that a typical firm in region r has technology

$$Y_{rt}(i) = A_{rt} \left(\frac{N_{rt}(i)}{1-\nu} \right)^{1-\nu} \left(\frac{M_{rt}(i)}{\nu} \right)^{\nu}, \tag{C.50}$$

where $M_{rt}(i)$ is a bundle of intermediate inputs composed of both domestic and foreign goods given by

$$M_{rt}(i) = \left((1-\upsilon_r^m)^{\frac{1}{\eta}} (M_{H,rt}(i))^{1-\frac{1}{\eta}} + (\upsilon_r^m)^{\frac{1}{\eta}} (M_{F,rt}(i))^{1-\frac{1}{\eta}} \right)^{\frac{\eta}{\eta-1}}, \tag{C.51}$$

where $M_{H,rt}(i)$ is an index of consumption of goods produced locally—that is, produced in region r—which is given by

$$M_{H,rt}(i) \equiv \left(\int_0^1 M_{H,rt}(i,j)^{\frac{\varepsilon-1}{\varepsilon}} dj \right)^{\frac{\varepsilon}{\varepsilon-1}}, \tag{C.52}$$

where $i, j \in [0,1]$ index varieties, $M_{H,rt}(i,j)$ is demand by firm i of intermediate inputs produced by firm j, and $M_{F,rt}(i)$ is demand by i for imported intermediate inputs.

Firm $i \in [0,1]$ in region r chooses labor and intermediate inputs to minimize total costs of production. Given our assumptions about technologies, the solution to the firm's cost minimization problem delivers the following nominal marginal cost of production:

$$NMC_{rt} = \frac{(W_{rt})^{1-\nu} (P_{M,rt})^{\nu}}{A_{rt}}, \tag{C.53}$$

where $P_{M,rt} = P_{rt}$ and P_{rt} is as in (12). We have omitted the index i from the nominal marginal cost, NMC_{rt}, because all firms face the same wage and prices of intermediate inputs. Hence, in the model with domestic and imported intermediate inputs, the real marginal cost in region r is given by

$$MC_{rt} = \frac{(W_{rt})^{1-\nu} (P_{rt})^{\nu}}{A_{rt} P_{H,rt}} = \frac{1}{A_{rt}} \left(\frac{W_{rt}}{P_{rt}} \right)^{1-\nu} \frac{P_{rt}}{P_{H,rt}}, \tag{C.54}$$

which, by rewriting $\frac{P_{rt}}{P_{H,rt}}$ as a function of the terms of trade, becomes

$$MC_{rt} = \frac{1}{A_{rt}} \left(\frac{W_{rt}}{P_{rt}} \right)^{1-\nu} \left((1-\upsilon_r) + \upsilon_r S_{rt}^{1-\eta} \right)^{\frac{1}{1-\eta}}. \tag{C.55}$$

Consider now the market clearing condition for tradable intermediates. We have that for each

$i \in [0,1]$,

$$\begin{aligned} Y_{rt}(i) &= C_{H,rt}(i) + \int_0^1 M_{H,rt}(j,i)\,dj + X_{rt}(i) \\ &= \left(\frac{P_{H,rt}(i)}{P_{H,rt}}\right)^{-\varepsilon}\left(C_{H,rt} + \int_0^1 M_{H,rt}(j)\,dj + X_{rt}\right). \end{aligned}$$

Aggregating over varieties we have that

$$\int_0^1 Y_{rt}(i)\,di = Y_{rt}\int_0^1 \left(\frac{P_{H,rt}(i)}{P_{H,rt}}\right)^{-\varepsilon}$$

where $Y_{rt} = C_{H,rt} + \int_0^1 M_{H,rt}(j)\,dj + X_{rt}$ and we can define the aggregate production function as

$$Y_{rt} \equiv \frac{\int_0^1 Y_{rt}(i)\,di}{\varrho_{rt}}, \tag{C.56}$$

where ϱ_{rt} is given by (B.40).

From the goods market clearing condition we have that

$$\begin{aligned} Y_{rt} &= C_{H,rt} + \int_0^1 M_{H,rt}(j)\,dj + X_{rt} \\ &= (1-\upsilon_r)\left(\frac{P_{H,rt}}{P_{rt}}\right)^{-\eta} C_{rt} + (1-\upsilon_r^m)\left(\frac{P_{H,rt}}{P_{rt}}\right)^{-\eta} M_{rt} + \frac{\upsilon^*}{I} S_{rt}^{\eta} Y_t^*. \end{aligned}$$

where $M_{rt} = \int_0^1 M_{rt}(i)\,di$. Following the same steps as with the model without intermediate inputs, we obtain that

$$\begin{aligned} V_{rt} &= (1-\upsilon_r)\left(\frac{P_{H,rt}}{P_{rt}}\right)^{-\eta} C_{rt}\left[(1-\upsilon_r) + \frac{\upsilon^*}{I}(Z_{rt})^{-\frac{1}{\sigma}} Q_{rt}^{\eta-\frac{1}{\sigma}}\right] \\ &= \left((1-\upsilon_r) + \upsilon_r S_{rt}^{1-\eta}\right)^{\frac{\eta}{1-\eta}} C_{rt}\left[(1-\upsilon_r) + \frac{\upsilon^*}{I}(Z_{rt})^{-\frac{1}{\sigma}} Q_{rt}^{\eta-\frac{1}{\sigma}}\right], \end{aligned} \tag{C.57}$$

where

$$V_{rt} \equiv Y_{rt} - (1-\upsilon_r^m)\left(\frac{P_{H,rt}}{P_{rt}}\right)^{-\eta} M_{rt}. \tag{C.58}$$

Hence, note that the right-hand-side of this equation is exactly the same as in the model without intermediate inputs. Therefore, log-linearizing this equation and following the steps in Appendix B, we obtain that the log of value added must be such that

$$v_{rt} = y_t^* + \frac{\phi_r}{\sigma} z_{rt} + \frac{(1-\upsilon_r) + \upsilon_r\omega_r}{\sigma} s_{rt}.$$

Consider now the optimal choice of intermediate inputs by each firm:

$$M_{rt}\frac{P_{rt}}{P_{H,rt}} = \nu MC_{rt}Y_{rt}. \tag{C.59}$$

This condition implies that

$$\begin{aligned}V_{rt} &= Y_{rt} - (1-\upsilon_r^m)\left(\frac{P_{H,rt}}{P_{rt}}\right)^{1-\eta}\nu MC_{rt}Y_{rt}\\ &= Y_{rt}\left(1-(1-\upsilon_r^m)\nu\left(\frac{P_{H,rt}}{P_{rt}}\right)^{1-\eta}MC_{rt}\right).\end{aligned}$$

Hence, we obtain the log-linear approximation of value added given by

$$\upsilon_{rt} = y_{rt} - (1-\upsilon_r^m)\nu\left(mc_{rt} - \upsilon_r s_{rt}\right), \tag{C.60}$$

where we have used again the fact that

$$\frac{P_{rt}}{P_{H,rt}} = \left((1-\upsilon_r) + \upsilon_r S_{rt}^{1-\eta}\right)^{\frac{1}{1-\eta}}.$$

Hence, conditions (C.59) and (C.60) imply that we can now express the terms of trade as a function of the output gap and marginal cost:

$$s_{rt} = \sigma_r^m\left(y_{rt} - y_t^*\right) - \sigma_r^m\nu(1-\upsilon_r^m)mc_{rt} - \sigma_r^m\frac{\phi_r}{\sigma}z_{rt}, \tag{C.61}$$

where

$$\sigma_r^m \equiv \frac{\sigma}{(1-\upsilon_r)+\upsilon_r\omega_r^m}, \tag{C.62}$$

and

$$\omega_r^m \equiv \omega_r - \sigma\nu(1-\upsilon_r^m). \tag{C.63}$$

Going back to the expression for real marginal cost, (C.54), we can use the intertemporal optimality condition, (13), the risk-sharing condition, (B.39), and labor demand to express the real wage as follows:

$$\begin{aligned}\frac{W_{rt}}{P_{rt}} &= \left(\left(Y_t^*\left(Z_{rt}\right)^{\frac{1}{\sigma}}\left(Q_{rt}\right)^{\frac{1}{\sigma}}\right)^{\sigma}\left((1-\nu)MC_{rt}\frac{P_{H,rt}}{P_{rt}}Y_{rt}\right)^{\varphi}\right)^{\frac{1}{1+\varphi}}\\ &= \left(\left(Y_t^*\left(Z_{rt}\right)^{\frac{1}{\sigma}}\left(Q_{rt}\right)^{\frac{1}{\sigma}}\right)^{\sigma}\left((1-\nu)MC_{rt}\left((1-\upsilon_r)+\upsilon_r S_{rt}^{1-\eta}\right)^{\frac{-1}{1-\eta}}Y_{rt}\right)^{\varphi}\right)^{\frac{1}{1+\varphi}}.\end{aligned}$$

We can now go back to the definition of real marginal cost in terms of real wages to obtain the following expression for the log of the real marginal cost:

$$mc_{rt} = \sigma y_t^* + \varphi y_{rt} + z_{rt} + \frac{1+\varphi}{1+\nu} a_{rt} + \upsilon_r s_{rt} + q_{rt}.$$

Using the fact that $q_{rt} = (1-\upsilon_r)s_{rt}$, we obtain that

$$mc_{rt} = \sigma y_t^* + y_{rt} + z_{rt} + \frac{1+\varphi}{1+\nu} a_{rt} + s_{rt},$$

which is equivalent to condition (B.47) adjusted for the use of intermediate inputs in production.

We can now use our expression for the terms of trade given by (C.61) to obtain the following expression for the real marginal cost:

$$\begin{aligned} mc_{rt} = & \frac{\sigma - \sigma_m^\upsilon}{1+\sigma_r^m \nu (1-\upsilon_r^m)} y_t^* + \frac{\varphi + \sigma_r^m}{1+\sigma_r^m \nu (1-\upsilon_r^m)} y_{rt} \\ & - \frac{1}{1+\sigma_r^m \nu (1-\upsilon_r^m)} \frac{1+\varphi}{1+\nu} a_{rt} + \frac{1-\sigma_m^\upsilon \frac{\phi_r}{\sigma}}{(1+\sigma_r^m \nu (1-\upsilon_r^m))} z_{rt}. \end{aligned}$$

Letting y_{rt}^n denote the natural level of log output, and following the same steps as for the case with no intermediate inputs in production, we obtain that log deviations from the natural real marginal cost are related to output deviations by

$$mc_{rt} - mc_r = \frac{\varphi + \sigma_r^m}{1+\sigma_r^m \nu (1-\upsilon_r^m)} \left(y_{rt} - y_{rt}^n \right).$$

As for the model with no intermediate inputs, note that openness only affects the slope of the Phillips curve by affecting the relation between real marginal cost and output. We now turn to expressing deviations in output as deviation in unemployment.

Combining labor demand with expression for real marginal cost as a function of real wages and relative prices, (C.54), and relative prices expressed as function of the terms of trade, we obtain that deviation in employment are given by

$$n_{rt} - n_{rt}^n = \frac{\nu}{1-\nu} \left(\upsilon_r (s_{rt} - s_{rt}^n) - (mc_{rt} - mc_r) \right) + (y_{rt} - y_{rt}^n).$$

Now, from (C.61) note that

$$s_{rt} - s_{rt}^n = \sigma_r^m (y_{rt} - y_{rt}^n) - \sigma_r^m \nu (mc_{rt} - mc_r),$$

which implies that

$$n_{rt} - n_{rt}^n = \left(1 + \frac{\nu}{1-\nu} \upsilon_r \sigma_r^m \right) (y_{rt} - y_{rt}^n) - \frac{\nu}{1-\nu} \left(1 + \upsilon_r \sigma_r^m \nu \right) (mc_{rt} - mc_r).$$

Given the relation between real marginal cost and the output gap, we obtain that

$$n_{rt} - n_{rt}^{n} = \frac{\kappa_r^m}{1+\sigma_r^m \nu (1-\upsilon_r^m)} \frac{1}{1-\nu} \left(y_{rt} - y_{rt}^{n}\right),$$

where

$$\iota_r^m \equiv \left(1+\nu(1-\upsilon_r^m)\sigma_r^m\right)\left(1+\nu\upsilon_r\sigma_r^m\right) - \nu\left(\left(1+\nu\upsilon_r\sigma_r^m\right)\left(\varphi+\sigma_r^m\right) + \left(1+\nu(1-\upsilon_r^m)\sigma_r^m\right)\right). \quad \text{(C.64)}$$

Therefore, we get that

$$mc_{rt} - mc_r = \frac{(1-\nu)(\varphi+\sigma_r^m)}{\iota_r^m} \left(n_{rt} - n_{rt}^{n}\right).$$

Note that if we consider the limiting case in which there are no intermediate inputs, $\nu \to 0$, then (C.64) collapses to the same equilibrium condition for the model without intermediate inputs. We turn now to how this condition changes for differing degrees of trade openness.

Note that now there are two different types of trade openness: openness in trade of final goods and openness in trade of intermediate inputs. Let us consider first the case of heterogeneity across regions in terms of their degree of openness with respect to final goods. Then, we have that

$$\frac{d}{d\upsilon_r} \frac{\varphi+\sigma_r^m}{\iota_r^m} = \frac{\frac{d\sigma_r^m}{d\upsilon_r}\iota_r^m - (\varphi+\sigma_r^m)\frac{d\iota_r^m}{d\upsilon_r}}{(\iota_r^m)^2}.$$

Considering the case in which $\sigma\eta > 1$, it can be readily seen that

$$\frac{d\sigma_r^m}{d\upsilon_r} = \frac{-\sigma \frac{d}{d\upsilon_r}[(1-\upsilon_r)+\upsilon_r\omega_r]}{((1-\upsilon_r)+\upsilon_r\omega_r^m)^2} < 0,$$

because of the conclusion from (C.50), but also because ω_r^m only depends on υ_r through ω_r, which enters linearly in (C.63). Hence, a sufficient condition for more open regions to have flatter Phillips curves is that $\frac{d\iota_r^m}{d\upsilon_r} > 0$. It can be shown that there are conditions for parameters under which this will be true, particularly for large values of φ.

Let us now turn to the case of heterogeneity across regions in terms of their degree of openness with respect to intermediate inputs. In this case, note that we also have that

$$\frac{d}{d\upsilon_r^m} \frac{\varphi+\sigma_r^m}{\iota_r^m} = \frac{\frac{d\sigma_r^m}{d\upsilon_r^m}\iota_r^m - (\varphi+\sigma_r^m)\frac{d\iota_r^m}{d\upsilon_r^m}}{(\iota_r^m)^2}.$$

It can be easily shown that

$$\frac{d\sigma_r^m}{d\upsilon_r^m} = -\left(\sigma_r^m\right)^2 \upsilon_r \upsilon < 0.$$

Hence, similarly to the previous case, a sufficient condition for more open regions to have flatter

Phillips curves is that $\frac{d\iota_r^m}{d\upsilon_r^m} > 0$. And it can be shown that there are reasonable parameter values restrictions under which the inequality holds.

Appendix D Model with Trade across Regions

Consider the case in which regions can trade with each other. We define aggregate consumption in region r as

$$C_{rt} \equiv \left((1-\upsilon_r) C_{L,rt}^{\frac{\eta-1}{\eta}} + \upsilon_r C_{F,rt}^{\frac{\eta-1}{\eta}} \right)^{\frac{\eta}{\eta-1}}$$

where

$$C_{L,rt} \equiv \left(\sum_{u=1}^{I} \left(\frac{1}{I}\right)^{\frac{1}{\gamma}} C_{ru,t}^{\frac{\gamma-1}{\gamma}} \right)^{\frac{\gamma}{\gamma-1}},$$

represents an aggregate of consumption produced across different regions in the home country and

$$C_{rut} = \left(\int_0^1 C_{rut}(i)^{\frac{\varepsilon-1}{\varepsilon}} di \right)^{\frac{\varepsilon}{\varepsilon-1}}.$$

is consumption of goods produced abroad. Hence, η parametrizes the trade elasticity for imports, while γ is related to the trade elasticity for trading goods across regions of the monetary union.[32]

In this case, we have that

$$P_{rt}^{1-\eta} = (1-\upsilon_r) P_{L,t}^{1-\eta} + \upsilon_r P_{F,t}^{1-\eta}$$

and

$$P_{L,rt}^{1-\gamma} = \sum_{u=1}^{I} \frac{1}{I} P_{H,ut}^{1-\gamma}.$$

For any region r, we can define the following price index and sourcing share associated to expenditure on domestic goods from other regions in the same country:

$$P_{R,rt}^{1-\gamma} \equiv \sum_{u \neq r} \frac{1}{I-1} P_{H,ut}^{1-\gamma}, \text{ and}$$

$$\Lambda_{L,rt} \equiv \frac{P_{H,rt} C_{rrt}}{P_{L,rt} C_{L,rt}}.$$

Then, following the same steps as in the proof of Lemma 2, we obtain that

$$(1+\pi_{rt})^{1-\eta} = (1+\pi_{L,rt})^{1-\eta} \left(\frac{1-\Lambda_{rt}}{1-\Lambda_{rt-1}} \right)^{-1}$$

[32] We have simplified the model by assuming that each region allocates total expenditure across all the regions of the monetary union equally, with shares given by $1/I$. This assumption can be easily generalized to arbitrary shares.

where

$$(1+\pi_{L,rt})^{1-\gamma} = (1+\pi_{H,rt})^{1-\eta}\left(\frac{1-\Lambda_{L,rt}}{1-\Lambda_{L,rt-1}}\right)^{-1}.$$

Therefore, in line with Lemma 2, a log-linear approximation of the previous equation delivers

$$\pi_{rt} = \pi_{H,rt} - \frac{d\Lambda_{L,rt}}{1-\gamma} - \frac{d\Lambda_{rt}}{1-\eta}. \tag{D.65}$$

In addition, following similar steps as in Appendix B, it can be shown that in this model

$$\sigma_r^{\upsilon} \equiv \frac{\sigma}{(1-\upsilon_r) + \upsilon_r\left[\sigma\eta + (1-\upsilon_r)(\sigma\eta - 1)\frac{1-\phi_{L,r}}{\upsilon_r}\right]},$$

where

$$\phi_{L,r} \equiv \frac{\frac{1}{I}\sum_{u=1}^{I}\frac{(1-\upsilon_u)^2}{(1-\upsilon_r)}}{\frac{1}{I}\sum_{u=1}^{I}(1-\upsilon_u) + \frac{\upsilon^*}{I}}.$$

Note that the previous expression simplifies to (B.42) when we assume that all regions are equally open. Similarly, if we assume no trade across regions. Hence, there are reasonable conditions on parameters that ensure that more open regions have flatter Phillips curves.

Note that according to (D.65), regional Phillips curves in this model incorporate the effects on inflation of buying cheaper goods from other regions of the monetary union and not only from abroad. Therefore, if there were historical data available for U.S. interstate trade at the aggregate or sectoral levels, we could construct a measure of penetration of domestic purchases from other regions in the monetary union and incorporate it into our estimation equation.

Appendix E Inflation and Unemployment Across U.S. States

Figure 6 shows average annual inflation across U.S. states from 1990 to 2017 using the data constructed by Hazell et al. (2022). The data shows sizable differences across state-specific non-housing inflation rates over this period of time. Figure 7 shows the variation in unemployment rates across U.S. states at the height of the Great Recession.

Figure 6: Non-Housing inflation across U.S. States: 1990-2016

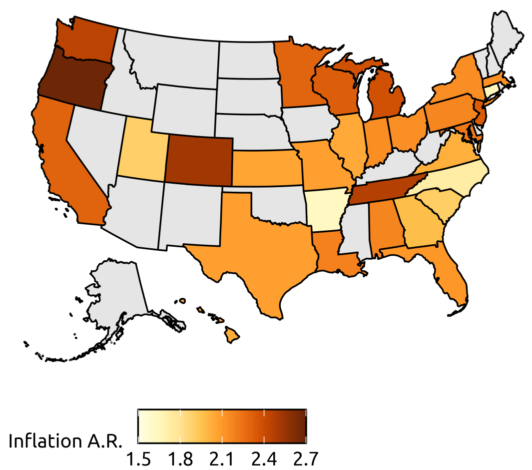
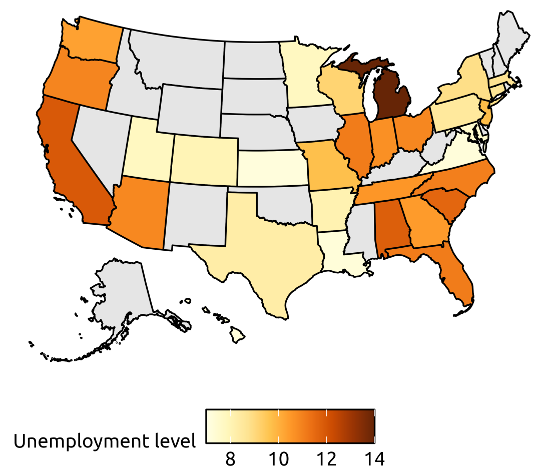


Figure 7: Unemployment rates across U.S. States: 2009



Notes: Data on prices at the state level sourced from Hazell et al. (2022). Unemployment rates constructed using data from the Quarterly Census of Employment and Wages (QCEW) by the BLS.