

Slicing the Pie: Quantifying the Aggregate and Distributional Effects of Trade

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We develop a multi-sector gravity model with heterogeneous workers to quantify the aggregate and group-level welfare effects of trade. The model generalizes the specific-factors intuition to a setting with labour reallocation, leads to a parsimonious formula for the group-level welfare effects from trade, and nests the aggregate results in Arkolakis, Costinot and Rodríguez-Clare (2012, “New Trade Models, Same Old Gains?”), *American Economic Review*, **102**, 94–130). We estimate the model using the structural relationship between China-shock driven changes in manufacturing employment and average earnings across US groups defined as commuting zones. We find that the China shock increases average welfare but some groups experience losses as high as four times the average gain. However, adjusting for plausible measures of inequality aversion barely affects the welfare gains. We also develop and estimate an extension of the model that endogenizes labour force participation and unemployment, finding similar welfare effects from the China shock.

Key words: International trade; Inequality, China shock; Sectoral reallocation; Local labour markets

JEL Codes: F11, F14, F16, F6, J24

1. INTRODUCTION

The recent empirical literature has made economists less sanguine about the overall benefits from increased trade integration. Although the notion that there are losers from trade is one of the oldest propositions in the field, recent empirical work exemplified most prominently by Autor, Dorn and Hanson (2013) has shown that the distributive implications of trade shocks in developed countries are stronger and more persistent than previously believed. In their survey of this work, Autor, Dorn and Hanson (2016) conclude that “it is incumbent on the literature to more convincingly estimate the gains from trade, such that the case for free trade is not based on the

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sway of theory alone, but on a foundation of evidence that illuminates who gains, who loses, by how much, and under what conditions.” In this article, we take a step in this direction—we develop and estimate a multi-sector gravity model of trade with heterogeneous labour and use it to quantify the group-level and aggregate welfare effects of the China shock and overall trade in the US.

Our baseline model combines three components: a multi-sector version of the [Eaton and Kortum \(2002\)](#) model as in [Costinot, Donaldson and Komunjer \(2012\)](#); a Roy model of the allocation of heterogeneous labour to sectors with a Fréchet distribution as in [Lagakos and Waugh \(2013\)](#); and the existence of different labour groups differing in their pattern of comparative advantage across sectors. The model yields a simple expression for the group-level welfare effects of trade that generalizes the formula previously shown by [Arkolakis, Costinot and Rodríguez-Clare \(2012\)](#) (henceforth ACR) to be valid for a wide class of gravity models. Compared to the ACR formula, ours has an extra term that captures the group-level effects of trade through changes in the vector of sector-specific wages. Thus, following a logic similar to that in the specific-factors model, groups with high employment shares in sectors that experience strong increases in import competition will fare worse than other groups. The strength of these distributional effects depends on the shape parameter of the Fréchet distribution, κ , which governs the degree of labour heterogeneity across sectors: if $\kappa \rightarrow 1$ then our model yields the same welfare implications as the one with sector-specific labour and distributional effects are strongest, while if $\kappa \rightarrow \infty$ then we are back to the single ACR formula applying to all groups.

Inspired by [Autor *et al.* \(2013\)](#) (henceforth ADH), our quantitative analysis focuses on the effect of the China shock on US workers grouped according to commuting zone. Not only is the focus on local labour markets important in its own right, but it also allows us to build on the empirical strategy developed by ADH to arrive at a credible estimate of κ . We employ an instrumental variable approach where the first stage estimates the group-level effect of the China shock on manufacturing employment, as in the reduced-form of one of the central regressions in ADH. The second stage then exploits the model-implied relationship between the projected change in the share of employment in non-manufacturing (one of the sectors in the model) and group-level average earnings. The estimation yields a value for κ around 1.5, which is in line with estimates of this Roy-Fréchet parameter in related contexts (e.g. [Hsieh, Hurst, Jones and Klenow, 2013](#); [Burstein, Morales and Vogel, 2019](#)).

Armed with our estimate of κ , we calibrate the China shock following a strategy similar to that in [Caliendo, Dvorkin and Parro \(2019\)](#) and then use the comparative-statics methodology in [Dekle, Eaton and Kortum \(2008\)](#) to compute the group-level and aggregate welfare effects of the China shock in the US. We find that a modest but non-negligible number of groups representing 13% of the population suffer welfare losses, and that those losses can be up to four times as high as the average gains. The welfare effects are spatially correlated, implying the existence of regions such as Southern Appalachia where most groups tend to experience low or negative effects. To compute the aggregate welfare effects of the trade shock, we ignore the possibility that losers are compensated and use a social welfare function with inequality aversion as in [Atkinson \(1970\)](#).¹ Since low-income commuting zones are actually less exposed to the China shock, the inequality-adjusted gains turn out to be slightly higher than gains without inequality aversion.

Moving beyond the China shock, we also use our model to compute the group-level and aggregate gains from trade, defined as in ACR as the negative of the losses from moving to

1. Recent papers that pursue a similar strategy in the trade context are [Antràs, De Gortari and Itskhoki \(2017\)](#), [Carrère, Grujovic and Robert-Nicoud \(2020\)](#), and [Artuc, Porto and Rijkers \(2019\)](#). [Antràs *et al.* \(2017\)](#) also consider the distortions associated with compensation and quantifies the associated effect on the gains from trade. While we do not address the issue of how to optimally compensate losers from trade in this article, our results on the substantial losses from trade for certain groups highlight the importance of this question for future research.

autarky. We again find that a small set of groups lose from trade, with one group experiencing losses of 4.8%, around three times the mean gain across all groups. Interestingly, however, the results imply that trade lowers inequality, and hence the inequality-adjusted gains from trade are slightly above those with no inequality aversion.

We consider a number of extensions to see how our baseline results change when allowing for tradable intermediate goods, trade costs within countries, imperfect substitutability of the labour input from skilled and unskilled labour within a commuting zone, and more disaggregation across groups so that they can vary by commuting zone, education, gender, and age. Tradable intermediates as in [Caliendo and Parro \(2015\)](#) amplify the gains from the China shock, so fewer groups experience losses, while allowing for trade costs across US states has the opposite effect. Imperfect substitutability between the labour input of college and non-college workers in each sector leads to an endogenous college premium (similar to the Heckscher–Ohlin model, although weakened by Roy heterogeneity), but this turns out not to have significant implications on the results for the welfare effects of the China shock in the baseline model, with most of those changes explained by commuting-zone rather than worker-type fixed effects. Finally, allowing groups to vary by education-by-gender or education-by-age within each commuting zone does not significantly affect the main conclusions derived in the baseline model—most importantly, the commuting zone to which a group belongs remains the main determinant of how it is affected by the China shock.

In a final extension, we introduce home production (as in [Caliendo *et al.* \(2019\)](#)) as well as search and matching frictions (as in [Kim and Vogel \(2021\)](#)), so that trade shocks now lead to endogenous employment changes both because of changes in labour-force participation as well as changes in involuntary unemployment. We estimate the full model again exploiting the instrumental-variables strategy inspired by ADH, but now taking into account the observed changes in labour-force participation and unemployment along with the observed changes in average earnings at the commuting-zone level. The model is now qualitatively and quantitatively consistent with observed employment changes, and yet the implications for welfare remain close to those in the baseline model.

Relative to the reduced-form approach in [Autor *et al.* \(2013\)](#), our general-equilibrium structural analysis enables us to compute the welfare gains and losses caused by the China shock across groups, rather than only the associated relative income effects. We can also quantify the welfare effects of counterfactual shocks such as a move to autarky or a decline in trade costs. Our framework thus serves to establish a formal connection between the fast-growing empirical literature on the distributional implications of trade shocks and the more theoretical approaches to compute aggregate welfare effects of trade surveyed in [Costinot and Rodríguez-Clare \(2014\)](#).

A growing body of empirical work documents substantial variation in local labour-market outcomes in response to national-level trade shocks. In addition to [Autor *et al.* \(2013\)](#), see for example [Dix-Carneiro and Kovak \(2017\)](#), [Kovak \(2013\)](#), and [Topalova \(2010\)](#).² Additionally, a large empirical and theoretical literature studies the distributional effects of trade—some important recent contributions are [Autor, Dorn, Hanson and Song \(2014\)](#), [Burstein and Vogel \(2017\)](#); [Costinot and Vogel \(2010\)](#), [Helpman, Itskhoki, Muendler and Redding \(2017\)](#), and [Krishna, Poole and Senses \(2012\)](#). A literature focusing specifically on the effect of trade shocks on the reallocation of workers across sectors finds significant effects for developed countries ([Revenga, 1992](#); [Artaç, Chaudhuri and McLaren, 2010](#), [Pierce and Schott, 2016](#)), although less so in developing countries (see e.g. [Goldberg and Pavcnik, 2007](#); [Dix-Carneiro, 2014](#)).

2. Other empirical papers exploring the effects of trade shocks on local labour markets are [Dauth, Findeisen and Suedekum \(2014\)](#); [Hakobyan and McLaren \(2016\)](#), and [Yi, Müller and Stegmaier \(2016\)](#).

Artuç *et al.* (2010), Dix-Carneiro (2014), and Adão (2016) also use a Roy model of the allocation of workers across sectors to offer a structural analysis of the distributional effects of trade shocks, but they focus on exogenous changes in the terms of trade in a small economy.³ We complement these papers by linking the Roy model of the labour market with a gravity model of trade and by using the resulting framework to provide a transparent way to quantify the aggregate and distributional welfare effects of trade.

Caliendo *et al.* (2019), Lee (2020), and Adão, Arkolakis and Esposito (2020) combine a gravity model of trade with a Roy model of labour allocation, as we do, but these papers focus on different questions: Caliendo *et al.* (2019) emphasize the dynamics of adjustment after an unexpected trade shock, Lee (2020) focuses on the implications for the skill premium, and Adão *et al.* (2020) centre on how the effect of the trade shock is affected by the interaction between workers' employment decisions and agglomeration economies at the local level.⁴ Relative to these papers, we derive an analytical expression for the group-level welfare effects of trade shocks that nests the ACR welfare formula and highlights the role of κ on the distributional effects of trade, and we introduce the concept of inequality-adjusted gains from trade to the gravity literature. On the empirical side, our article provides a link between the reduced-form results of ADH and the estimation of κ that is needed to compute the group-level welfare effects of trade.⁵

The rest of this article is structured as follows. Section 2 describes the baseline model and presents our theoretical results. The data are described in Section 3, and Section 4 discusses the structural estimation of the model. Section 5 presents the results of the calibrated China shock for welfare of US groups, while Section 6 computes the aggregate and group-level gains from trade. Sections 7 and 8 present several extensions of the baseline model, and Section 9 offers some concluding thoughts.

2. THEORY

We present a multi-sector, multi-country, Ricardian model of trade with heterogeneous workers. There are N countries and S sectors. Each sector is modelled as in Eaton and Kortum (2002)—henceforth EK; there is a continuum of goods, preferences across goods within a sector s are CES with elasticity of substitution σ_s , and technologies have constant returns to scale with productivities that are distributed Fréchet with shape parameter $\theta_s > \sigma_s - 1$ and level parameters T_{is} in country i and sector s . Preferences across sectors are Cobb–Douglas with shares β_{is} . There are iceberg trade costs $\tau_{ijs} \geq 1$ to export goods in sector s from country i to country j , with $\tau_{iis} = 1$.

On the labour side, we assume that there are G_i groups of workers in country i . A worker from group g in country i (henceforth simply group ig) has a number of efficiency units z_s in sector s drawn from a Fréchet distribution with shape parameter $\kappa_{ig} > 1$ and scale parameters A_{igs} . Thus, workers within each group are *ex ante* identical but *ex post* heterogeneous due to different ability

3. Other structural analyses of trade liberalization and labour market adjustments are Coşar (2013), Coşar, Guner and Tybout (2016), Kambourov (2009), and Kim and Vogel (2021). There is also a literature on the impact of trade on poverty and the income distribution using a Computable General Equilibrium (CGE) methodology—see for example Cockburn, Decaluwé and Robichaud (2008).

4. While all the papers cited so far focus on the differential impact of trade through the earnings channel, another set of papers focuses on the expenditure channel—see Atkin and Donaldson (2015), Faber (2014), Fajgelbaum and Khandelwal (2016), and Porto (2006). More recent contributions by Borusyak and Jaravel (2018) and Artuc *et al.* (2019) consider both channels simultaneously.

5. Our article is also related to Hsieh and Ossa (2016), who use a gravity framework to conduct a comparative-static analysis in the style of Dekle *et al.* (2008) to quantify the aggregate effects of the China shock, and to Amiti, Dai, Feenstra and Romalis (2020) and Bai and Stumpner (2019), both of which estimate the effect of the China shock on the US consumer price index.

draws across sectors, as in Roy (1951), while workers across groups also differ in that they draw their abilities from different distributions. The number of workers in a group is fixed and denoted by L_{ig} . In Section 8 we extend the model to allow for non-employment and unemployment by introducing home production and search-and-matching frictions, respectively.

If $\kappa_{ig} \rightarrow \infty$ for all ig and $A_{igs} = 1$ for all igs , the model collapses to the multi-sector EK model developed in Costinot *et al.* (2012), while if $\kappa_{ig} \rightarrow 1$ for all ig then the model has the same welfare and counterfactual implications as the model in which labour is sector specific.⁶ On the other hand, if $\tau_{ijs} \rightarrow \infty$ for all $j \neq i$ and $G_i = 1$ then economy i is in autarky and collapses to the Roy model in Lagakos and Waugh (2013) (see also Hsieh *et al.*, 2013).⁷

2.1. Equilibrium

To determine the equilibrium of the model, it is useful to separate the analysis into two parts: the determination of labour demand in each sector in each country as a function of wages, which comes from the EK part of the model; and the determination of labour supply to each sector in each country as a function of wages, which comes from the Roy part of the model.

Since workers are heterogeneous in their sector productivities, the supply of labour to each sector is upward sloping, and hence wages can differ across sectors. However, since technologies and goods prices are national, wages cannot differ across groups. Let wages per efficiency unit in sector s of country i be denoted by w_{is} . From EK, we know that the demand for efficiency units in sector s in country i is

$$\frac{1}{w_{is}} \sum_j \lambda_{ijs} \beta_{js} X_j,$$

where X_j is total expenditure by country j and λ_{ijs} are sectoral trade shares given by

$$\lambda_{ijs} = \frac{T_{is} (\tau_{ijs} w_{is})^{-\theta_s}}{\sum_l T_{ls} (\tau_{ljs} w_{ls})^{-\theta_s}}. \tag{1}$$

For future purposes, also note that the price index in sector s in country j is

$$P_{js} = \zeta_s^{-1} \left(\sum_i T_{is} (\tau_{ijs} w_{is})^{-\theta_s} \right)^{-1/\theta_s}, \tag{2}$$

where $\zeta_s \equiv \Gamma(1 - \frac{\sigma_s - 1}{\theta_s})^{1/(1 - \sigma_s)}$.⁸

6. The only difference between the model with sector-specific labour and ours with $\kappa_{ig} \rightarrow 1$ is that in ours the elasticity of labour supply to any particular sector with respect to the wage in that sector goes to one and not zero. However, for $\kappa_{ig} \rightarrow 1$ the reallocation of workers across sectors has no effect on the relative supply of efficiency units of labour across sectors—see Equation (4). Note that $\kappa_{ig} \rightarrow 1$ implies that mean efficiency units per worker goes to infinity—when we report results for this limit we are implicitly normalizing efficiency units by $\Gamma(1 - 1/\kappa_{ig})$, where $\Gamma()$ is the Gamma function.

7. There are two sources of comparative advantage in this model: first, as in Costinot *et al.* (2012), differences in T_{is} drive sector-level (Ricardian) comparative advantage; second, differences in A_{igs} lead to factor-endowment driven comparative advantage. Given the nature of our comparative statics exercise, however, the source of comparative advantage will not matter for the results—only the actual sector-level specialization as revealed by the trade data will be relevant.

8. As shown in ACR, a multi-sector version of the Armington model would be a workable substitute for the EK-side of the model. The Krugman (1980) model or the Melitz (2003) model with a Pareto distribution (as in Chaney, 2008) would also work, though these models would introduce extra terms because of entry effects—see Costinot and Rodríguez-Clare (2014) and Kucheryavyi, Lyn and Rodríguez-Clare (2018).

Labour supply is determined by workers' choices regarding which sector to work in. Let $\mathbf{z} = (z_1, z_2, \dots, z_S)$ and let $\Omega_{is} \equiv \{\mathbf{z} \text{ s.t. } w_{is}z_s \geq w_{ik}z_k \text{ for all } k\}$. A worker with productivity vector \mathbf{z} in country i will apply to sector s iff $\mathbf{z} \in \Omega_{is}$. Let $F_{ig}(\mathbf{z})$ be the joint probability distribution of \mathbf{z} for workers of group ig . From [Lagakos and Waugh \(2013\)](#) and [Hsieh et al. \(2013\)](#), we know that the share of workers in group ig that apply to sector s is

$$\pi_{igs} \equiv \int_{\Omega_{is}} dF_{ig}(\mathbf{z}) = \frac{A_{igs}w_{is}^{\kappa_{ig}}}{\Phi_{ig}^{\kappa_{ig}}}, \quad (3)$$

where $\Phi_{ig}^{\kappa_{ig}} \equiv \sum_k A_{igk}w_{ik}^{\kappa_{ig}}$.⁹ Under the assumption that efficiency units from different workers are perfect substitutes in production, we just care about the sum of efficiency units supplied to a sector among all workers in a group. For group ig and sector s , this is

$$\mathcal{Z}_{igs} \equiv L_{ig} \int_{\Omega_{is}} z_s dF_{ig}(\mathbf{z}) = \xi_{ig} \frac{\Phi_{ig}}{w_{is}} \pi_{igs} L_{ig}, \quad (4)$$

where $\xi_{ig} \equiv \Gamma(1 - 1/\kappa_{ig})$. One implication of this result is that even if wages per efficiency unit of labour w_{is} differ across sectors, expected income per worker is equalized. That is, for each group ig and for all s we have

$$\frac{w_{is}\mathcal{Z}_{igs}}{\pi_{igs}L_{ig}} = \xi_{ig}\Phi_{ig}.$$

This is a special implication of the Fréchet distribution and it implies that the share of income obtained by workers of group ig in sector s (i.e. $w_{is}\mathcal{Z}_{igs}/\sum_k w_{ik}\mathcal{Z}_{igk}$) is also given by π_{igs} . Note also that total labour income in group ig is $Y_{ig} \equiv \sum_s w_{is}\mathcal{Z}_{igs} = \xi_{ig}\Phi_{ig}L_{ig}$, while total labour income in country i is $Y_i \equiv \sum_{g \in G_i} Y_{ig}$.

Allowing for trade imbalances D_j via transfers as in [Dekle et al. \(2008\)](#), we have

$$X_j = Y_j + D_j, \quad (5)$$

with $\sum_j D_j = 0$. Finally, combining the supply and demand sides of the economy, the excess demand for efficiency units in sector s of country i is

$$ELD_{is} \equiv \frac{1}{w_{is}} \sum_j \lambda_{ijs} \beta_{js} X_j - \sum_{g \in G_i} \mathcal{Z}_{igs}. \quad (6)$$

Since λ_{ijs} , Y_j , and \mathcal{Z}_{igs} are functions of the whole matrix of wages $\mathbf{w} \equiv \{w_{is}\}$, the system $ELD_{is} = 0$ for all i and s is a system of equations in \mathbf{w} whose solution gives the equilibrium wages given some choice of numeraire.

2.2. Comparative statics

Consider some change in trade costs or technology parameters. We proceed as in [Dekle et al. \(2008\)](#) and solve for the proportional change in the endogenous variables. Formally, using notation

9. This result and the ones below generalize easily to a setting with correlation in workers' ability draws across sectors. In this case, the dispersion parameter κ_{ig} is replaced by $\kappa_{ig}/(1 - \rho_{ig})$, where ρ_{ig} measures the correlation parameter of ability draws across sectors for each worker.

$\hat{x} \equiv x'/x$, we consider shocks $\hat{\tau}_{ijs}$ for $i \neq j$, \hat{D}_j , \hat{A}_{igs} and \hat{T}_{is} . The counterfactual equilibrium entails $ELD'_{is} = 0$ for all i, s . Noting that $w'_{is} Z'_{igs} = \hat{\pi}_{igs} \hat{Y}_{ig} \pi_{igs} Y_{ig}$, equation $ELD'_{is} = 0$ can be written as

$$\sum_j \hat{\lambda}_{ijs} \lambda_{ijs} \beta_{js} \left(\sum_{g \in G_j} \hat{Y}_{jg} Y_{jg} + \hat{D}_j D_j \right) = \sum_{g \in G_i} \hat{\pi}_{igs} \hat{Y}_{ig} \pi_{igs} Y_{ig} \tag{7}$$

with

$$\hat{Y}_{ig} = \left(\sum_k \pi_{igk} \hat{A}_{igk} \hat{w}_{ik}^{\kappa_{ig}} \right)^{1/\kappa_{ig}}, \tag{8}$$

$$\hat{\lambda}_{ijs} = \frac{\hat{T}_{is} (\hat{\tau}_{ijs} \hat{w}_{is})^{-\theta_s}}{\sum_k \lambda_{kjs} \hat{T}_{ks} (\hat{\tau}_{kjs} \hat{w}_{ks})^{-\theta_s}}, \tag{9}$$

and

$$\hat{\pi}_{igs} = \frac{\hat{A}_{igs} \hat{w}_{is}^{\kappa_{ig}}}{\sum_k \pi_{igk} \hat{A}_{igk} \hat{w}_{ik}^{\kappa_{ig}}}. \tag{10}$$

Given values for parameters θ_s and κ_{ig} ; data on income levels, Y_{ig} , trade imbalances, D_j , trade shares, λ_{ijs} , expenditure shares, β_{is} , labour allocation shares π_{igs} , and labour endowments, L_{ig} ; and the shocks to trade costs, $\hat{\tau}_{ijs}$, trade imbalances, \hat{D}_j , and productivity levels, \hat{A}_{igs} and \hat{T}_{is} , we can solve for changes in wages, \hat{w}_{is} , from the system of equations associated with (7)–(10), and then solve for all other relevant changes, including changes in trade shares using (9) and changes in employment shares using (10).

2.3. Group-level welfare effects

Our measure of welfare of individuals in group ig is *ex ante* real income, $W_{ig} \equiv \frac{Y_{ig}/L_{ig}}{P_i}$.¹⁰ We are interested in the change in W_{ig} caused by a shock to trade costs or foreign technology levels, henceforth simply referred to as a “foreign shock.” Cobb–Douglas preferences imply that $P_i = \prod_s P_{is}^{\beta_{is}}$, and hence

$$\hat{W}_{ig} = \hat{Y}_{ig} \prod_s \hat{P}_{is}^{-\beta_{is}}. \tag{11}$$

From (2) and (9) and given $\hat{T}_{is} = 1$ for all s in domestic country i , we have $\hat{P}_{is} = \hat{w}_{is} \hat{\lambda}_{iis}^{1/\theta_s}$ while from (8) and (10) we have $\hat{Y}_{ig} = \hat{w}_{is} \hat{\pi}_{igs}^{-1/\kappa_{ig}}$. Combining these two results with (11) we arrive at the following proposition:

Proposition 1. *Given some shock to trade costs or foreign technology levels, the percentage change in the real wage of group g in country i is given by*

$$\hat{W}_{ig} = \prod_s \hat{\lambda}_{iis}^{-\beta_{is}/\theta_s} \cdot \prod_s \hat{\pi}_{igs}^{-\beta_{is}/\kappa_{ig}}. \tag{12}$$

10. This is the same as utility if there were no trade imbalances. In the presence of trade imbalances, utility would instead be $(1+d_{ig})W_{ig}$, where $d_{ig} \equiv D_{ig}/Y_{ig}$ and D_{ig} is the trade deficit of group ig . The formulas below would need to be adjusted to capture changes in d_{ig} by multiplying by $\frac{1+d_{ig}d'_{ig}}{1+d_{ig}}$. Since we do not know how a country’s trade imbalance is allocated to groups, we do not observe d_{ig} . Our approach in the quantitative analysis will be to first use the model to shut down trade imbalances and then use the resulting data for our quantitative analysis.

The RHS of the expression in (12) has two components: $\prod_s \hat{\lambda}_{iis}^{-\beta_{is}/\theta_s}$ and $\prod_s \hat{\pi}_{igs}^{-\beta_{is}/\kappa_{ig}}$, with all variation across groups coming from the second term. If $\kappa_{ig} \rightarrow \infty$ for all $g \in G_i$ then the gains for all groups in country i are equal to $\prod_s \hat{\lambda}_{iis}^{-\beta_{is}/\theta_s}$, which is the multi-sector formula for the welfare effect of a trade shock in ACR. It is easy to show that the term $\prod_s \hat{\lambda}_{iis}^{-\beta_{is}/\theta_s}$ corresponds to the change in real income given wages while the term $\prod_s \hat{\pi}_{igs}^{-\beta_{is}/\kappa_{ig}}$ corresponds to the change in real income for group ig coming exclusively from changes in wages \hat{w}_{is} for $s = 1, \dots, S$.¹¹

The term $\prod_s \hat{\pi}_{igs}^{-\beta_{is}/\kappa_{ig}}$ is related to the change in the degree of specialization of group ig . We can use the Kullback–Leibler (KL) divergence as a way to define the degree of specialization of a group. Formally, the KL divergence from $\boldsymbol{\pi}_{ig} \equiv \{\pi_{ig1}, \pi_{ig2}, \dots, \pi_{igS}\}$ to $\boldsymbol{\beta}_i \equiv \{\beta_{i1}, \beta_{i2}, \dots, \beta_{iS}\}$ is given by

$$D_{KL}(\boldsymbol{\beta}_i \parallel \boldsymbol{\pi}_{ig}) \equiv \sum_s \beta_{is} \ln(\beta_{is}/\pi_{igs}).$$

If group ig was in group-level autarky (i.e. not trading with any other group or country) then $\pi_{igs} = \beta_{is}$ for all s . Thus, $D_{KL}(\boldsymbol{\beta}_i \parallel \boldsymbol{\pi}_{ig})$ is a measure of the degree of specialization as reflected in the divergence from the actual distribution, $\boldsymbol{\pi}_{ig}$, to what it would be in autarky, $\boldsymbol{\beta}_i$.¹² We can now write

$$\prod_s \hat{\pi}_{igs}^{-\beta_{is}/\kappa_{ig}} = \exp\left(\frac{1}{\kappa_{ig}} \left[D_{KL}(\boldsymbol{\beta}_i \parallel \boldsymbol{\pi}'_{ig}) - D_{KL}(\boldsymbol{\beta}_i \parallel \boldsymbol{\pi}_{ig}) \right]\right). \quad (13)$$

This implies that, apart from the common term $\prod_s \hat{\lambda}_{iis}^{-\beta_{is}/\theta_s}$, the welfare effect of a trade shock on a particular group in country i is determined by the change in the degree of specialization of that group as measured by the KL divergence, multiplied by the degree of heterogeneity in worker productivity across sectors as captured by $1/\kappa_{ig}$. For example, a group with high employment in textiles would become less specialized and gain less (or even lose) from trade if a foreign shock leads the country to import disproportionately more textiles. On the other hand, groups specialized in exporting sectors gain more from trade than the country as a whole.

Of course, Proposition 1 cannot in general be used to go from observables and elasticities to welfare. We first need to use the model to compute $\{\hat{\lambda}_{iis}\}$ and $\{\hat{\pi}_{igs}\}$ for whatever shock we are interested in. This is also true in ACR, where the formula $\hat{W}_i = \hat{\lambda}_{ii}^{-\theta}$ is only directly applicable to find the gains from trade relative to autarky. Still, we highlight the formula as Proposition 1 because it shows clearly how our model extends the results in ACR, and because it is informative about the way in which the model works, in particular by pointing out the role of the elasticities and the changes in trade and employment shares. In Section 2.6 below, we show an approximate

11. The result in Proposition 1 can alternatively be derived by first applying the envelope theorem to the consumption and labour allocation problem at the group level,

$$d \ln W_{ig} = \sum_s \pi_{igs} d \ln w_{js} - \sum_{i,s} \beta_{js} \lambda_{ijs} d \ln(w_{is} \tau_{ijs}).$$

We can then proceed as in ACR to substitute for $d \ln w_{js}$ and $d \ln(w_{is} \tau_{ijs})$ in this expression. From the trade side of the model, we have $\frac{d \ln(\lambda_{ijs}/\lambda_{jjs})}{d \ln(w_{is} \tau_{ijs}/w_{js})} = -\theta_s$, while from the labour side we have $\frac{d \ln(\pi_{jgs}/\pi_{jkk})}{d \ln(w_{js}/w_{jk})} = -\kappa_{jg}$. Solving for $d \ln(w_{is} \tau_{ijs})$ and $d \ln w_{js}$ from these two equations, respectively, and then plugging back into the expression for $d \ln W_{ig}$ above yields $d \ln W_{ig} = -\sum_s \beta_{js} \left[\frac{d \ln \pi_{jgs}}{\kappa_{jg}} + \frac{d \ln \lambda_{ijs}}{\theta_s} \right]$. Integration leads to the result in (12).

12. The KL divergence was introduced by Kullback and Leibler (1951) and is also known as relative entropy. The KL divergence from q to p , $D_{KL}(p|q) = \sum_i p_i \ln(p_i/q_i)$, is equal to the difference between the cross entropy from q to p , $H(p, q) = -\sum_i p_i \ln(q_i)$, and the entropy of p , $H(p) = -\sum_i p_i \ln(p_i)$, $D_{KL}(p|q) = H(p, q) - H(p)$. Subtracting $H(q)$ ensures that $D_{KL}(p|p) = 0$.

formula that uses observables and elasticities to compute the effect of trade on a group’s relative income.¹³

2.4. Aggregate welfare effects

We define aggregate welfare as aggregate real income, $W_i \equiv Y_i/P_i$.¹⁴ The aggregate welfare effect can be obtained from Proposition 1 as $\hat{W}_i = \hat{Y}_i/\hat{P}_i = \sum_{g \in G_i} (Y_{ig}/Y_i) \hat{W}_{ig}$, leading to

$$\hat{W}_i = \prod_s \hat{\lambda}_{iis}^{-\beta_{is}/\theta_s} \cdot \sum_{g \in G_i} \left(\frac{Y_{ig}}{Y_i} \right) \prod_s \hat{\pi}_{igs}^{-\beta_{is}/\kappa_{ig}}. \tag{14}$$

The welfare effect of a trade shock is no longer given by the multi-sector ACR term (i.e. $\hat{W}_i \neq \prod_s \hat{\lambda}_{iis}^{-\beta_{is}/\theta_s}$). This is because a trade shock will affect sector-level wages w_{is} , and this in turn will affect welfare through its impact on income and sector-level prices.

2.5. Aggregate and group-level gains from trade

Following ACR, we define the gains from trade as the negative of the proportional change in real income for a shock that takes the economy back to autarky: $GT_i \equiv 1 - \hat{W}_i^A$ and $GT_{ig} \equiv 1 - \hat{W}_{ig}^A$.

A move to autarky for country i entails $\hat{\tau}_{ijs} = \infty$ for all s and all $i \neq j$ and $\hat{D}_i = 0$. Conveniently, solving for changes in wages in country i (\hat{w}_{is} for $s = 1, \dots, S$) from Equation (7) only requires knowing the values of employment shares, income levels and expenditure shares for country i , namely β_{is} for all s , Y_{ig} for all g , and π_{igs} for all g, s . This can be seen by letting $\hat{\tau}_{ijs} \rightarrow \infty$ in Equation (7), which yields

$$\beta_{is} \sum_{g \in G_i} \hat{Y}_{ig} Y_{ig} = \sum_{g \in G_i} \hat{\pi}_{igs} \hat{Y}_{ig} \pi_{igs} Y_{ig}. \tag{15}$$

Let $r_{is} \equiv \sum_{g \in G_i} \pi_{igs} Y_{ig}/Y_i$ be the share of sector s in total output in country i and note that country i engages in inter-industry trade as long as $r_{is} \neq \beta_{is}$ for some s .

Proposition 2. Assume that $\kappa_{ig} = \kappa_i$ for all $g \in G_i$. If $\kappa_i < \infty$ and country i engages in inter-industry trade, then the aggregate gains from trade are strictly higher than those that arise in the limit as $\kappa_i \rightarrow \infty$.

Supplementary Appendix B.1 has the proof. To understand this result, it is useful to consider the simpler case with a single group of workers, $G_i = 1$. In this case, a move back to autarky would

13. We comment briefly on how our model relates to the one in ADH. They derive their regression equations from a log-linear approximation of the equilibrium conditions of a multi-sector gravity model of trade with homogeneous and perfectly mobile workers across sectors, but with each group modelled as a separate economy. In this case, all the variation in the effects of a shock across groups arises because of different terms of trade effects. In our baseline model, technologies are national and there are no trade costs among groups within countries, so with homogenous and perfectly mobile workers across sectors the terms of trade effects would be the same for all groups. Instead, worker heterogeneity implies that some groups of workers are more closely attached to some sectors, and it is this that generates variation in the effect of trade shocks across groups.

14. Aggregate real income gives the utility level of all agents if income is shared equally among all individuals in the economy. Also, if there is no risk sharing, W_i gives utility “behind the veil of ignorance” if agents are risk neutral, as further explained in Section 2.7.

imply

$$\hat{W}_i^A = \prod_s \lambda_{iis}^{\beta_{is}/\theta_s} \cdot \exp \left[-\frac{1}{\kappa_i} D_{KL}(\boldsymbol{\beta}_i \parallel \mathbf{r}_i) \right].$$

If there is inter-industry trade then $D_{KL}(\boldsymbol{\beta}_i \parallel \mathbf{r}_i) > 0$ so (given \mathbf{r}_i) a finite κ_i implies a lower \hat{W}_i^A than in the multi-sector ACR formula. Intuitively, a finite κ_i introduces more “curvature” to the PPF, making it harder for the economy to adjust as it moves to autarky. This implies higher losses if the economy were to move to autarky, and hence higher gains from trade. Proposition 2 establishes that this result generalizes to the case $G_i > 1$.

Turning to the group-specific gains from trade, we again use the KL measure of specialization to understand whether a group gains more or less than the economy as a whole. The results of the previous section imply that the gains from trade for group ig are

$$GT_{ig} = 1 - \prod_s \lambda_{iis}^{\beta_{is}/\theta_s} \cdot \exp \left(\frac{1}{\kappa_{ig}} \left[D_{KL}(\boldsymbol{\beta}_i \parallel \boldsymbol{\pi}_{ig}^A) - D_{KL}(\boldsymbol{\beta}_i \parallel \boldsymbol{\pi}_{ig}) \right] \right).$$

The term $D_{KL}(\boldsymbol{\beta}_i \parallel \boldsymbol{\pi}_{ig}^A) - D_{KL}(\boldsymbol{\beta}_i \parallel \boldsymbol{\pi}_{ig})$ could be positive or negative, depending on whether group ig becomes more or less specialized with trade as measured by the KL divergence.

Consider a group ig that happens to have efficiency parameters $(A_{ig1}, \dots, A_{igS})$ that give it a strong comparative advantage in a sector s for which the country as a whole has a comparative disadvantage, as reflected in positive net imports in that sector. Group ig would be highly specialized in s when the country is in autarky but that specialization would diminish as the country starts trading with the rest of the world. As a consequence, the KL degree of specialization falls with trade for group ig , implying lower gains relative to other groups in the economy.

2.6. A Bartik approximation

Focusing on the implications of a foreign shock on a group’s relative income, equation (8) implies that

$$\frac{\hat{Y}_{ig}}{\hat{Y}_i} = \left(\sum_s \pi_{igs} \left(\frac{\hat{w}_{is}}{\hat{Y}_i} \right)^{\kappa_{ig}} \right)^{1/\kappa_{ig}}. \quad (16)$$

Since wages are not observable, it is convenient to derive an approximation for this expression that uses changes in output shares, \hat{r}_{is} rather than \hat{w}_{is} . Assuming that $\kappa_{ig} = \kappa_i$ for all $g \in G_i$ and recalling that $r_{is} \equiv \sum_{g \in G_i} \pi_{igs} Y_{ig}/Y_i$, equations (8) and (10) imply:

$$\hat{r}_{is} = \left(\frac{\hat{w}_{is}}{\hat{Y}_i} \right)^{\kappa_i} \sum_{g \in G_i} \frac{(Y_{ig}/Y_i) \pi_{igs}}{r_{is}} \left(\frac{\hat{Y}_{ig}}{\hat{Y}_i} \right)^{1-\kappa_i}.$$

The term $\frac{(Y_{ig}/Y_i) \pi_{igs}}{r_{is}}$ captures group ig ’s share of country i ’s total output of sector s , and $(\hat{Y}_{ig}/\hat{Y}_i)^{1-\kappa_i}$ is an adjustment to take into account how $(\hat{Y}_{ig}/\hat{Y}_i) \pi_{igs}$ deviates from $(\hat{w}_{is}/\hat{Y}_i)^{\kappa_i}$ for group ig . The sum on the RHS of the previous equation is then an overall adjustment for how \hat{r}_{is} may deviate from $(\hat{w}_{is}/\hat{Y}_i)^{\kappa_i}$. For κ close to 1 or for shocks that do not lead to large differences in \hat{Y}_{ig}/\hat{Y}_i from 1 for groups with large weights in sector s , that adjustment will be small, and $\hat{r}_{ik} \approx (\hat{w}_{is}/\hat{Y}_i)^{\kappa_i}$, so Equation 16 yields

$$\frac{\hat{Y}_{ig}}{\hat{Y}_i} \approx \left(\sum_k \pi_{igk} \hat{r}_{ik} \right)^{1/\kappa_i}. \quad (17)$$

In the quantitative analysis in Sections 5 and 6, we will see that this equation provides a very good approximation of the model implied group-level relative income effects of the China shock and the move back to autarky for the US. The benefit of this result is that \hat{r}_{is} is observable in the data. Thus, if we can identify the impact of a foreign shock on output shares, then we can use this Bartik-style result to compute approximate relative income changes across groups.

This result is particularly useful for the shock that takes country i back to autarky. For that case, we have $\hat{r}_{is} = \beta_{is}/r_{is}$, and hence, we obtain an approximate sufficient statistic for a group's gains from trade relative to the aggregate gains:

$$\frac{\hat{Y}_{ig}^A}{\hat{Y}_i^A} \approx I_{ig}^{1/\kappa_i} \equiv \left(\sum_s \pi_{igs} \frac{\beta_{is}}{r_{is}} \right)^{1/\kappa_i}. \tag{18}$$

We can think of β_{is}/r_{is} as an index of the degree of import competition in industry s and I_{ig} as an index of import competition faced by group g . Thus, for a move back to autarky, the change in relative income levels across groups is approximated by the index of import competition that we can directly observe in the data elevated to the power $1/\kappa_i$. Since a foreign shock does not affect the autarky equilibrium, we can also use the result in (18) to rewrite the approximation in (17) for any foreign shock in terms of the change in the index of import competition, $\frac{\hat{Y}_{ig}}{\hat{Y}_i} \approx \hat{I}_{ig}^{-1/\kappa_i}$.¹⁵

2.7. Inequality-adjusted welfare effects

We follow Atkinson (1970) and think about social welfare as a (geometric) average of welfare across all individuals with a constant inequality aversion parameter $\rho > 0$ (with $\rho \neq 1$ to simplify the exposition below). Since the z_s for workers in group ig is distributed Fréchet with scale parameter A_{igs} and shape parameter κ_{ig} , then income $\max_s w_{is} z_s$ for workers in group ig is distributed Fréchet with scale parameter $\Phi_{ig}^{\kappa_{ig}}$ and shape parameter κ_{ig} . Social welfare in country i is then

$$U_i = \frac{1}{P_i} \left(\sum_{g \in G_i} \int_0^\infty y^{1-\rho} l_{ig} dH_{ig}(y) \right)^{\frac{1}{1-\rho}},$$

with $H_{ig}(y) = \exp(-\Phi_{ig}^{\kappa_{ig}} y^{-\kappa_{ig}})$. Integrating and assuming that $\kappa_{ig} = \kappa_i$ yields

$$U_i = \tilde{\xi}_i \left(\sum_g l_{ig} W_{ig}^{1-\rho} \right)^{\frac{1}{1-\rho}}, \tag{19}$$

where $l_{ig} \equiv L_{ig}/L_i$ and $\tilde{\xi}_i \equiv \frac{\Gamma(1-\frac{1-\rho}{\kappa_i})^{\frac{1}{1-\rho}}}{\Gamma(1-\frac{1}{\kappa_i})}$.

The inequality-adjusted welfare effect of a foreign shock is defined as $\hat{U}_i - 1$ whereas the inequality-adjusted gains from trade are defined as $IGT_i \equiv 1 - \hat{U}_i^A$. If $\rho = 0$ then these

15. This result can also be derived directly from (17) by noting that $\sum_k \pi_{igk} \hat{r}_{ik} = \frac{\sum_s \pi_{igs} \beta_{is}/r_{is}}{\sum_s \pi_{igs} \beta_{is}/r'_{is}} = 1/\hat{I}_{ig}$. See [Supplementary Appendix B.2](#).

measures correspond to those defined above, namely $\hat{W}_i - 1$ and $GT_i \equiv 1 - \hat{W}_i^A$.¹⁶ To write these results in terms of observables and the endogenous group-level welfare changes \hat{W}_{ig} , let $\omega_{ig} \equiv \frac{l_{ig}(Y_{ig}/L_{ig})^{1-\rho}}{\sum_h l_{ih}(Y_{ih}/L_{ih})^{1-\rho}}$ be a modified weight for group ig in country i welfare that appropriately accounts for the social value of income accruing to groups with different income levels. Then, simple algebra reveals that

$$\hat{U}_i = \left(\sum_g \omega_{ig} \hat{W}_{ig}^{1-\rho} \right)^{\frac{1}{1-\rho}}. \quad (20)$$

2.8. Extensions

The combination of a stylized model of the labour market with a standard multi-sector gravity model delivers clean analytical results on the group-level welfare effects of trade shocks, while nesting the ACR welfare formula. The implied distributional effects are closely approximated by Bartik-style changes in import competition and can be integrated into an aggregate measure of inequality-adjusted welfare effects. We will show below that this baseline model is sufficient to provide a structural framework for the empirical analysis of changes in import-competition on group-level changes in income and unemployment, and in our counterfactual analysis we will document how the Roy component of our model leads to strong distributional effects of a prominent trade shock, the China shock. In Sections 7–9, we will study several extensions of the baseline model that allow for variation across groups not only by commuting zone but also by gender, age and education, imperfect substitutability between skilled and unskilled labour, endogenous employment levels, tradable intermediate goods, and trade costs within the US, studying for each case the associated implications.¹⁷

3. DATA

For our quantitative analysis, we define groups based on geographic location. We follow ADH in using commuting zones (CZs) as geographic units to define local labour markets.¹⁸ This leaves us with a total of 722 groups (CZs). All countries other than the US are assumed to have a single group.

Since our baseline estimation follows ADH as closely as possible, we employ the same data sources and definitions. These include labour income and employment status from the American Community Survey (ACS) and decennial censuses, employment shares across industries for each

16. A Rawlsian approach to social welfare entails $\rho \rightarrow \infty$ and $\hat{U}_i = \min_g W'_{ig} / \min_g W_{ig}$. If $\arg \min_g W'_{ig} = \arg \min_g W_{ig} = h$ then $\hat{U}_i = \hat{W}_{ih}$, but of course this need not be the case. We discuss plausible values for ρ in Section 5.

17. In the [Supplementary Appendix Section J](#), we also develop an extension with mobility of workers across groups, motivated by the case in which groups correspond to commuting zones. Given the lack of the necessary data, we have not explored the quantitative implications of this extension.

18. Our assumption of fixed groups applied to this setting implies no mobility across local labour markets. We view this as a reasonable assumption in light of existing literature that finds little evidence of trade exposure causing population shifts across local labour markets. See, for example, ADH for the US, [Dauth et al. \(2014\)](#) for Germany, and [Dix-Carneiro and Kovak \(2017\)](#) for Brazil.

commuting zone from the County Business Patterns database (CBP), and trade flows from the UN Comtrade database.¹⁹ As in ADH, we focus our analysis on the periods 1990–2000 and 2000–7.²⁰

Due to data limitations, our simulation analysis is restricted to the time period 2000–7 and uses aggregated industry definitions. Our choice of time horizon (2000–7) resulted from the data requirement on bilateral trade flows from the World Input-Output Database (WIOD), which are only available starting 1995.²¹ We chose to have more aggregated sectors in order to link the labour data with WIOD figures in a consistent manner. These aggregated sectors, listed in Appendix Table A.1, are based on the 1987 SIC classification codes. We aggregate all manufacturing industries into 13 sectors which roughly correspond to two-digit ISIC Rev. 3 codes. The remaining sectors, excluding public administration and the non-profit sector, are aggregated to one non-manufacturing sector.²²

Supplementary Appendix C describes in detail the construction of our dataset and the definition of our variables. It also details the supplementary data employed in our model extensions and robustness tests.²³

4. ESTIMATION

The κ parameter is central to our model as it jointly affects the aggregate and the distributional effects from trade. In this section, we propose and then implement an estimation strategy for this parameter that builds on the seminal work of ADH, and in particular on their findings that across commuting zones, the China shock leads to a significant contraction in manufacturing employment and a decline in earnings.

4.1. From model to regression equation

We now restrict attention to the US and therefore drop the country subscript. We estimate a common value for κ_g across groups and hence impose $\kappa = \kappa_g$. From equations (8) and (10), we then obtain $\hat{y}_g = \hat{A}_{gs}^{1/\kappa} \hat{w}_s \hat{\pi}_{gs}^{-1/\kappa}$, where $y_g \equiv Y_g/L_g$ is defined as average income per employed worker in group g . This expression holds for any sector s , and says that, conditional on \hat{w}_s and \hat{A}_{gs} , $\hat{\pi}_{gs}^{-1/\kappa}$ serves as a sufficient statistic for the change in a group’s income. Intuitively, given \hat{w}_s and \hat{A}_{gs} , then $\hat{\pi}_{gs} > 1$ ($\hat{\pi}_{gs} < 1$) implies that wages (or productivity shocks) weighted by employment shares in other sectors must have been negative (positive) for group g , leading workers in that

19. In all our estimations, we follow very closely the data construction, sample restrictions, variable definitions and model specifications of ADH. We obtain data from the decennial censuses and ACS from IPUMS USA (Ruggles, Flood, Goeken, Grover, Meyer, Pacas and Sobek, 2020). We then apply the same sample restrictions and use the same industry classification (3-digit SIC codes), same set of covariates, etc. For a detailed description of our data, see Supplementary Appendix C.

20. As in ADH, we make adjustments to the data in order to put the two periods on a comparable decadal scale. For the period 2000–7, we multiply employment, income, and trade changes with a factor of 10/7. Since trade figures are only available from 1991 for the time period 1991–2000, we multiplied trade growth with the factor 10/9.

21. The World Input–Output Database (WIOD) is discussed in Timmer, Dietzenbacher, Los, Stehrer and Vries (2015).

22. Since we require consistency between the trade and labour data, for US groups we first set $Y_{igs} = \frac{\pi_{igs}^{CBP} Y_{ig}^{CBP}}{\sum_h \pi_{ih}^{CBP} Y_{ih}^{CBP}} Y_{is}^{WIOD}$, where the superscript denotes the data source, and then focus on π_{igs} as shares of earnings, $\pi_{igs} = \frac{Y_{igs}}{\sum_k Y_{igk}}$. Recall that in our Roy–Fréchet framework the share of workers of any group ig in sector s is the same as the share of earnings derived from working in that sector.

23. These include data on unemployment, home production and alternative group definitions employed in the extensions presented in Sections 7 and 8.

group to move to (out of) sector s . The parameter κ determines how large the loss in relative income is for a given $\hat{\pi}_{gs}$.²⁴

Applying this expression to the non-manufacturing sector, $s = NM$, adding a t subscript to denote time periods, and taking logs yields

$$\ln \hat{y}_{gt} = \delta_t + \beta \ln \hat{\pi}_{gNMt} + \varepsilon_{gt}, \quad (21)$$

where $\delta_t \equiv \ln \hat{w}_{NMt}$, $\beta \equiv -1/\kappa$ and $\varepsilon_{gt} \equiv \ln \hat{A}_{gNMt}^{1/\kappa}$. We can use this equation to estimate κ from a cross-group regression of $\ln \hat{y}_{gt}$ on $\ln \hat{\pi}_{gNMt}$ (pooling across periods), instrumented as in ADH as explained below.²⁵ Focusing on the non-manufacturing sector allows us to build on the primary finding in ADH, namely the contraction in manufacturing employment caused by the China shock, and leads to a stronger first stage in our IV estimation.

4.2. Empirical strategy

The model implies that the regressor is correlated with the error term in the regression equation (21), i.e., $\mathbb{E}[\ln \hat{\pi}_{gNMt} \cdot \varepsilon_{gt}] \neq 0$. Hence, instead of running a simple OLS regression, we pursue an instrumental-variable strategy to obtain consistent estimates, using the exact same China shock variable as constructed by ADH. Specifically, the instrumental variable we use is

$$Z_{gt} \equiv \sum_{s \in M} \pi_{gst-10} \Delta IP_{st}^{\text{China} \rightarrow \text{Other}}, \quad (22)$$

where M refers to the subset of manufacturing sub-industries and

$$\Delta IP_{st}^{\text{China} \rightarrow \text{Other}} \equiv \frac{\Delta \text{Imports}_{st}^{\text{China} \rightarrow \text{Other}}}{L_{st-10}^{\text{US}}},$$

where L_{st-10}^{US} denotes US employment in sector s in year $t-10$, $\text{Imports}_{st}^{\text{China} \rightarrow \text{Other}}$ are imports from China by a group of countries similar to the US, and Δ refers to the change over period t .²⁶ We focus on the two time periods used in ADH, namely 1990–2000 and 2000–7. In the construction of instrument Z_{gt} , the π_{gst-10} are measured for 397 manufacturing sub-industries, employing data from the County Business Patterns.²⁷

24. One way to understand why, conditional on \hat{w}_s and \hat{A}_{gs} , $\hat{\pi}_{gs}^{-1/\kappa}$ serves as a sufficient statistic for \hat{y}_g is as follows. For any sector s , we know that $\frac{\hat{\pi}_{gk}}{\hat{\pi}_{gs}} = \left(\frac{\hat{w}_k}{\hat{w}_s}\right)^\kappa$ for all g and k and $\sum_k \pi_{gk} \hat{\pi}_{gk} = 1$ for all g , hence $\hat{\pi}_{gs} \sum_k \pi_{gk} \left(\frac{\hat{w}_k}{\hat{w}_s}\right)^\kappa = 1$ for all g . This implies that groups more exposed to relative wage declines will have a higher $\hat{\pi}_{gs}$, implying that $\hat{\pi}_{gs}$ acts as a sufficient statistic for such exposure and the associated income change. The same reasoning implies that groups with higher employment shares in sectors experiencing relative wage declines will have higher expansions in sectors that originally had lower employment shares, leading to a larger decline in specialization as measured by the KL divergence and a relative fall in income.

25. The absence of within-country trade costs is an important assumption in the derivation of this estimation equation, as it ensures that $\ln \hat{w}_{NMt}$ does not vary across commuting zones. We examine the sensitivity of our estimation and simulation results to this assumption in Section 7.3.

26. The use of countries similar to the US is meant to proxy for changes in sectoral import-competition from China in the US. This set of countries is identical to the set in ADH and consists of Australia, Denmark, Finland, Germany, Japan, and Spain, Switzerland and New Zealand. Countries are selected based on having a similar income level as the US, but direct neighbours are excluded.

27. Although here we have a higher level of disaggregation than in the quantitative analysis below (where we use only 13 manufacturing sectors), it is still the case that $\sum_s \pi_{gst-10}$ equals the total share of employment in manufacturing.

Our estimation of κ from (21) with $\hat{\pi}_{gNMt}$ instrumented by Z_{gt} is consistent if the instrument is relevant, $\text{cov}(Z_{gt}, \ln \hat{\pi}_{gNMt}) \neq 0$, and satisfies the exclusion restriction, $\text{cov}(Z_{gt}, \varepsilon_{gt}) = 0$, where these covariances are taken with respect to g for each t .²⁸ Regarding the first condition, large enough technology shocks in manufacturing sectors in China, $\hat{T}_{\text{China},st}$ for $s \in M$, would increase Chinese exports to other countries and to the US, leading to the contraction of the manufacturing sector in the most exposed groups and implying that $\text{cov}(Z_{gt}, \ln \hat{\pi}_{gNMt}) > 0$, as found by ADH and confirmed below. In turn, a sufficient condition for the exclusion restriction to hold is that the shocks $\{\hat{A}_{gNMt}\}$ be mean independent of group g lagged employment shares, $\mathbb{E}(\varepsilon_{gt} | \pi_{gt-10}) = 0$, as this would immediately imply that $\text{cov}(Z_{gt}, \varepsilon_{gt}) = 0$.²⁹

The condition that the shocks $\{\hat{A}_{gNM}\}$ be mean independent of group g lagged employment shares would fail, for example, if non-manufacturing productivity tended to fall in groups with high employment shares in unskilled-intensive manufacturing sectors. This would create a negative correlation between the instrument and the residual and lead to a downward bias in $\hat{\kappa}$. Similar concerns led ADH to add a set of commuting-zone variables as controls in their estimation. Such controls can be accommodated in our model by assuming that the unobserved productivity shocks are correlated with a vector of group-level variables X_{gt} . Formally, assuming that $\varepsilon_{gt} = X'_{gt} \Theta + \epsilon_{gt}$, where Θ is a vector of parameters, then the estimating equation becomes

$$\ln \hat{y}_{gt} = \delta_t + \beta \ln \hat{\pi}_{gNMt} + X'_{gt} \Theta + \epsilon_{gt}. \tag{23}$$

We use the same set of control variables as in ADH.³⁰ The sufficient condition for the exclusion restriction to hold is now weaker, as we need $\mathbb{E}(\epsilon_{gt} | \pi_{gt-10}, X_{gt}) = 0$ rather than $\mathbb{E}(\varepsilon_{gt} | \pi_{gt-10}) = 0$.

If condition $\mathbb{E}(\epsilon_{gt} | \pi_{gt-10}, X_{gt}) = 0$ holds, then any Bartik-type instrument combining employment shares and sector level changes would satisfy the exclusion restriction, even if it were correlated to US sector-level supply or demand shocks. This reveals an important difference between our paper and ADH: whereas the goal in ADH was to identify the causal effect of the China shock on income and employment in the US, our goal is instead to estimate parameter κ . ADH needed to avoid confounding the China shock with US import demand shocks, but in our case those import demand shocks can be part of the variation used to identify κ under the condition $\mathbb{E}(\epsilon_{gt} | \pi_{gt-10}, X_{gt}) = 0$.

Accordingly, below we consider two alternative instruments, one using the change in exports by China to the US rather than to other countries, $Z_{gt} \equiv \sum_{s \in M} \pi_{gst} \Delta IP_{st}^{\text{China} \rightarrow \text{US}}$, and the other

28. To understand what these conditions entail, think about the model as the data generating process: given initial data, parameters $\{\theta_s\}$ and κ , and a set of exogenous shocks $\{\hat{A}_{igs}\}$ and $\{\hat{T}_{is}\}$, the model in hat changes in Equations (7)–(10) generates $\{\hat{y}_{ig}\}$, $\{\hat{\pi}_{igNM}\}$ and $\{\Delta IP_s^{\text{China} \rightarrow \text{Other}}\}$ (for each period). We think of our data as $\{\Delta IP_s^{\text{China} \rightarrow \text{Other}}\}$ and a subsample \hat{y}_g and $\hat{\pi}_{gNM}$ for $g = 1, \dots, \tilde{G} < G$ (again focusing on the US and suppressing the country subindex). Consistency is for the limit as $\tilde{G} \rightarrow \infty$.

29. In principle, we could also apply estimation equation (21) to each of our 13 manufacturing sectors. Having selected some manufacturing sector s as the basis for the regression (instead of non-manufacturing), we would then construct an instrument as above but leaving out that sector. It turns out that this instrument lacks power regardless of the manufacturing sector s that we consider: changes in imports in manufacturing sectors $s' \neq s$ do not provide a good instrument for the change in the share of employment in manufacturing sector s . This stands in contrast to how changes in imports in all manufacturing sectors affect changes in the share of employment in non-manufacturing, which was one of the key results in ADH.

30. These controls are lagged manufacturing shares, Census division fixed effects, and beginning-of-period conditions (% college educated, % foreign-born, % employment among women, % employment in routine occupations, and the average offshorability index).

using the simple Bartik expression $Z_{gt} \equiv \ln \sum_s \pi_{gst} \hat{r}_{st}$, where r_{st} is the share of sector s in total sales in year t .^{31,32}

Although $\mathbb{E}(\epsilon_{gt} | \boldsymbol{\pi}_{gt-10}, \mathbf{X}_{gt}) = 0$ is a sufficient condition for identification, it is by no means a trivial assumption—see [Goldsmith-Pinkham, Sorkin and Swift \(2020\)](#), [Borusyak, Hull and Jaravel \(2022\)](#), henceforth BHJ, provide an alternative condition for instrument validity by focusing on the exogeneity of the shocks (rather than the shares), and thinking of consistency in terms of the number of sectors rather than the number of groups. This condition is

$$\text{cov}(Z_{gt}, \epsilon_{gt}) = \sum_{s \in M} \Delta IP_{st}^{\text{China} \rightarrow \text{Other}} \mathbb{E}[\pi_{gst-10} \epsilon_{gt}] \rightarrow 0$$

as the number of sectors in M goes to infinity. The sector shares π_{gst-10} are now allowed to be correlated with the error term ϵ_{gt} , as long as this correlation is orthogonal to the sector-specific China shocks. This alternative condition to instrument validity has implications for the computation of standard errors as well as additional specification and over-identification tests, which we discuss below when we present the estimation results.

4.3. Estimation results

Table 1 presents the results of the IV regression described above, with slight variations in the construction of the instrument.³³ The first row shows our second-stage results, while the third row has the corresponding estimate $\hat{\kappa} = -1/\hat{\beta}$, and the fourth row displays the F-statistic from the first stage. The first-stage F-statistics are always sufficiently high, which is not surprising given the central finding in ADH on the contraction of manufacturing due to the China shock. Most importantly, our estimated values for $\hat{\kappa}$ range from 1.42 to 2.79, and these estimates are statistically significant.³⁴

Our range of estimated values for κ is consistent with estimates of supply elasticities obtained by [Hsieh et al. \(2013\)](#) and [Burstein et al. \(2019\)](#), across occupations. Despite different modelling and estimation approaches, these papers find parameters of productivity dispersion (analogous to our κ) between 1.2 and 3.44.

Our baseline identifying assumption (the exogeneity of employment shares) is not directly testable. We can, however, identify which industry shares are driving our results, and frame our assumptions in terms of those industries. Following [Goldsmith-Pinkham et al. \(2020\)](#) (whose assumptions are analogous to ours), we computed Rotemberg weights for each industry

31. Due to data limitations, we use contemporaneous shares π_{igt} (instead of lagged shares) when constructing the instrument based on Chinese imports to the US. Using lagged shares requires detailed 1980 data, which is difficult to obtain. These instruments correspond to the endogenous regressor in ADH, which is the main source of our data. Identification in this case relies on the assumption that $\mathbb{E}(\epsilon_{gt} | \boldsymbol{\pi}_{gt}, \mathbf{X}_{gt}) = 0$.

32. The assumption that $\mathbb{E}(\epsilon_{gt} | \boldsymbol{\pi}_{gt-10}, \mathbf{X}_{gt}) = 0$ also implies that conventional inference methods are valid in our setting. Since this condition can be derived directly from our model's structural assumptions, we present conventional standard errors in our primary results (clustering at the state level to remain consistent with the results in ADH).

33. Reassuringly, the estimates line up reasonably well across the different columns. We performed a standard Hansen-J overidentification test which fails to reject that the four estimates are statistically the same (our Hansen-J statistic has a p -value of 0.346).

34. Our estimation strategy relies on assuming a Fréchet distribution, which restricts the mechanisms through which the China shock affects inequality (see [Adão, 2016](#)). In particular, the Fréchet assumption implies that there will be no effect of the China shock on within-group inequality. In [Supplementary Appendix Table D.1](#), we empirically test whether this is the case. We do so by running reduced form regressions with different measures of within-group inequality as the dependent variables and China shock measures as regressors of interest. The majority of our estimates yield no statistically significant evidence that the China shock increased within-group inequality.

TABLE 1
Estimation of κ

	(1)	(2)	(3)	(4)
	$\ln \hat{y}_g$	$\ln \hat{y}_g$	$\ln \hat{y}_g$	$\ln \hat{y}_g$
$\ln \hat{\pi}_{NM}$	-0.358* (0.211)	-0.639** (0.303)	-0.704** (0.295)	-0.487** (0.183)
Implied κ	2.79 (1.643)	1.56 (0.742)	1.42 (0.594)	2.05 (0.773)
F-first stage	58.46	24.02	29.52	67.87
Observations	1,444	1,444	1,444	1,444
R^2	0.683	0.667	0.662	0.677
Import penetration	Other (lagged)	Other (no lag)	US (no lag)	US Bartik

Notes: IV-estimation results for specification (23), where y_g is average earnings per worker, and π_{gNM} is the employment share in non-manufacturing, measured using the CBP data. The columns differ in the construction of the instrument: column (1) uses the exact instrument borrowed from ADH $Z_{gt} \equiv \sum_{s \in M} \pi_{gst-10} \Delta IP_{st}^{China \rightarrow Other}$, column (2) uses $Z_{gt} \equiv \sum_{s \in M} \pi_{gst} \Delta IP_{st}^{China \rightarrow Other}$, column (3) uses $Z_{gt} \equiv \sum_{s \in M} \pi_{gst} \Delta IP_{st}^{China \rightarrow US}$, and column (4) uses our Bartik variable for the US: $Z_{gt} \equiv \ln \sum_s \pi_{gst} \hat{r}_{st}$. Due to data constraints on π_{gst-10} , we have not constructed $Z_{gt} \equiv \sum_{s \in M} \pi_{gst-10} \Delta IP_{st}^{China \rightarrow US}$. Standard errors are clustered at the state level and reported in parentheses, with * $p < 0.1$, ** $p < 0.05$, *** $p < 0.001$. The first row shows the second-stage results, while the third row has the corresponding κ estimates implied by the model and the fifth row displays the F-statistic from the first stage. All regressions include the same controls employed in ADH's preferred specification: lagged manufacturing shares, period fixed effects, Census division fixed effects, and beginning-of-period conditions (% college educated, % foreign-born, % employment among women, % employment in routine occupations, and the average offshorability index).

share and found that the China shock instrument is driven by a few industries such as Electronic Computers (see [Supplementary Appendix Tables D.2 and D.4](#)). We can then view our orthogonality assumption as stating that CZs with high shares in these industries did not experience better or worse local productivity shocks to the non-manufacturing sector. To alleviate concerns of exogeneity violations for important industries, we constructed different versions of the China shock in a manner inspired by ADH, namely by sequentially leaving each of the top five sectors out. Reassuringly, the estimates for κ do not change significantly and all fall within the range of 1.17 to 3.12 (see [Supplementary Appendix Tables D.3 and D.5](#)).

If assumption $\mathbb{E}(\epsilon_{gt} | \pi_{gt-10}, X_{gt}) = 0$ is violated and identification relies instead on the alternative assumption provided by BHJ, then standard inference methods might not be appropriate³⁵. In [Supplementary Appendix D](#), we present our baseline point estimates of [Table 1](#) with standard errors computed as proposed by BHJ (see column 1 of [Supplementary Table D.6](#)). As expected, the standard errors become slightly larger, but our estimates retain conventional levels of statistical significance. In [Supplementary Appendix D](#), we also describe additional specification tests suggested by BHJ, and overall find that while results are less precise they nevertheless remain consistent with our baseline findings (see [Supplementary Table D.6](#)). We also find supporting evidence for the orthogonality of the shocks from an over-identification test (see [Supplementary Table D.7](#)).

35. As pointed out by [Adão, Kolesár and Morales \(2019\)](#), this is especially true in cases where the error terms are correlated for groups with similar employment shares. In fact the BHJ-corrected standard errors are asymptotically equivalent to those derived by [Adão et al. \(2019\)](#). We present the estimated standard errors in brackets in [Supplementary Appendix Table D.6](#). We were not able to compute the [Adão et al. \(2019\)](#) standard errors for the instruments constructed using beginning-of-period shares. This was due to the shares not satisfying the necessary rank conditions (a problem that does not apply to the lagged shares employed in their own analysis). This problem has been recently documented by BHJ, who point out that under certain rank conditions their inference approach is feasible whereas [Adão et al. \(2019\)](#)'s is not.

For the next section, where we will run simulations to analyse the quantitative role of κ in our framework, we will set our preferred value at $\kappa = 1.5$. In addition, we will also show results for $\kappa \rightarrow 1$ (the theoretical lower bound for κ), and for $\kappa = 3$ (twice our preferred value).

5. AGGREGATE AND DISTRIBUTIONAL EFFECTS OF THE RISE OF CHINA

While existing research (e.g. ADH) has found strong distributional implications of the “rise of China” across local labour markets in the US, this empirical research remains largely silent on the associated group-level and aggregate welfare effects. We now perform counterfactual simulations with our model to shed light on this question.³⁶

5.1. Calibrating the China shock

We model the rise of China as sector-specific technology shocks, $\hat{T}_{\text{China},s}$. We calibrate these shocks such that for each sector, the simulated changes in US expenditure shares on Chinese goods match the change in these expenditure shares that is driven by the rise of China.³⁷ The first step is to obtain predicted changes in US expenditure shares from running a specification similar to ADH’s first-stage regression,

$$\hat{\lambda}_{\text{China,US},s} = \alpha + \beta \hat{\lambda}_{\text{China,Other},s} + \varepsilon_s,$$

where $\hat{\lambda}_{\text{China,Other},s} \equiv \frac{\sum_{j \in \text{Other}} \lambda_{\text{China},j,s}^{2007}}{\sum_{j \in \text{Other}} \lambda_{\text{China},j,s}^{2000}}$. In a second step, we calibrate the technology shocks $\hat{T}_{\text{China},s}$ so that the model-implied changes in the US expenditure shares on imports from China, $\hat{\lambda}_{\text{China,US},s}$, match the predicted values from the first step.

5.2. Aggregate and distributional welfare effects

The results for the US welfare effects of the China shock as calibrated above are shown in Table 2 for four different values of κ : 1, 1.5, 3 and ∞ , and for $\theta_s = 5$ for all s .³⁸ The third row shows standard errors for each statistic based on the estimated $\kappa = 1.5$ in Table 1.³⁹ The first column shows the aggregate welfare effect for the case with no inequality aversion, \hat{W}_{US} , while the next four columns show the mean, the coefficient of variation (CV), and the minimum and maximum for the group-level welfare changes, $\hat{W}_{\text{US},g}$. The last column shows the welfare effect according to the multi-sector ACR formula.

36. In all the ensuing counterfactual exercises, we follow Head and Mayer (2014) and set $\theta_s = 5$ for all s . We perform our counterfactual exercises on data without trade deficits, which we obtain by first simulating the trade equilibrium with balanced trade. This preliminary simulation is always performed with our preferred value of $\kappa = 1.5$.

37. This calibration is inspired by the procedure in Caliendo *et al.* (2019), who calibrate $\hat{T}_{\text{China},s}$ to match predicted changes in US imports from China. Instead of imports, we focus on the expenditure shares $\lambda_{\text{China,US},s}$, and thereby avoid any complications arising from matching sectoral deflators for US imports across simulations and data.

38. To more clearly see the impact of κ on the welfare effects from the China shock, the results for different values of κ correspond to the shock $\hat{T}_{\text{China},s}$ as calibrated for $\kappa = 1.5$. Separately calibrating $\hat{T}_{\text{China},s}$ for each value of κ leads to broadly similar results—see Supplementary Appendix Table E.1.

39. These standard errors are computed based on the delta method. Each statistic of interest is a function $f(\hat{\beta})$ of our estimated $\hat{\beta}$, and so we compute its standard error as $SE(f(\hat{\beta})) = SE(\hat{\beta})|f'(\hat{\beta})|$, with $SE(\hat{\beta}) = 0.303$ as in column 2 in Table 1, and $f'(\hat{\beta})$ being the numerical derivative computed using simulations.

TABLE 2
The welfare effects of the China shock on the US

κ	\widehat{W}_{US}	Mean	CV	Min.	Max.	$\prod_s \widehat{\lambda}_s^{-\beta_s/\theta_s}$
$\rightarrow 1$	0.24	0.32	1.24	-1.44	2.30	0.14
1.5	0.22	0.29	1.01	-1.18	1.65	0.15
	(0.02)	(0.03)	(0.23)	(0.3)	(0.6)	(0.01)
3.0	0.20	0.26	0.70	-0.74	0.96	0.16
$\rightarrow \infty$	0.20	0.20	0	0.20	0.20	0.20

Notes: The first column displays the aggregate welfare effect of the China shock for the US, in percentage terms $100(\widehat{W}_{US} - 1)$, and the second column shows the mean welfare effect: $100(\frac{1}{G} \sum_g \widehat{W}_{US,g} - 1)$. The third column shows the coefficient of variation (CV), and for the fourth and fifth column, we have $\text{Min.} \equiv \min_g 100(\widehat{W}_{US,g} - 1)$ and $\text{Max.} \equiv \max_g 100(\widehat{W}_{US,g} - 1)$, respectively. The final column displays the multi-sector ACR term $100\left(\prod_s \widehat{\lambda}_{US,US,s}^{-\beta_{US,s}/\theta_s} - 1\right)$. The values for $\widehat{T}_{China,s}$ are calibrated for $\kappa = 1.5$. The third row has standard errors in parentheses, computed using the delta method and numerical derivatives with respect to $\widehat{\beta} = 1/\widehat{\kappa}$, for each statistic when $\kappa = 1.5$. We provide robustness checks for these numbers in Supplementary Appendix Tables E.1 and E.2.

Focusing first on the results for our preferred value of $\kappa = 1.5$, the model implies US aggregate welfare gains from the rise of China of 0.22%, with an average gain across groups of 0.29%.⁴⁰ The CV is 1.01, and the range is [-1.18%, 1.65%], implying a maximum loss that is four times the average gain.⁴¹ While 15 groups lose more than 0.5% of their real income, 109 groups gain more than 0.5% of their income. In total, 87% of groups, also representing 87% of the population, experience positive gains from the rise of China (see Appendix Figure A.1b).⁴²

There is a strong geographical correlation in the gains and losses from the China shock, as is clear from Figure 1, which plots the geographical distribution of the welfare effects from this shock. In the Eastern half of the country, largely excluding the coastal commuting zones, many groups experience below median gains. Particularly in the North East and in Central and Southern Appalachia, there is a strong concentration of commuting zones in the bottom third of the gains distribution.⁴³

The distributional impact of the China shock depends on κ , as a lower κ leads to higher dispersion in the gains from trade due to a stronger pattern of worker-level comparative advantage. The simulation results confirm this theoretical prediction, as both the CV and the difference between maximal and minimal $\widehat{W}_{US,g}$ tend to zero as κ approaches infinity (see Table 2). For $\kappa \rightarrow 1$, the CV reaches a maximum at 1.24, and the range is [-1.44%, 2.3%]. Table 2 also shows that for $\kappa \leq 3$ there are groups who lose substantially from the rise of China.⁴⁴

40. To provide context for this number, Hsieh and Ossa (2016) find welfare gains for the US between 0 and 0.03%. The difference with our results is likely due to the fact that we calibrate Chinese technology growth to fit predicted Chinese exports, whereas Hsieh and Ossa (2016) calculate technological growth based on firm-level data.

41. The standard errors shown in the table reveal that the aggregate and average results are estimated quite precisely, while this is less so for the results capturing the dispersion of the welfare effects. For example, the maximum loss would be up to 1.87% at the 95% confidence level rather than the estimated 1.18%. To some extent, we will see how this matters for the inequality-adjusted gains from trade by looking at the case with $\kappa = 1$.

42. Of course, when a commuting zone experiences positive gains, this does not imply that all workers in that group gain. For instance, workers who stay in a shrinking sector may lose real income. Importantly though, the focus of our model is on group-level average changes in income, not on tracking income changes at the individual level.

43. Our quantitative analysis assumes that the effect of the China shock on prices is the same across groups. This is consistent with Bai and Stumpner (2019), who find “no evidence for heterogeneous effects across consumer groups by income or region.”

44. Appendix Figures A.1 and A.2 visualize how the distributional impact of the China shock diminishes as κ increases, by plotting the full distribution of $\widehat{W}_{US,g}$ for different values of κ . To further understand the role of κ , recall that equation (13) shows how a higher κ directly mitigates the distributional impact of any reallocation, while Appendix

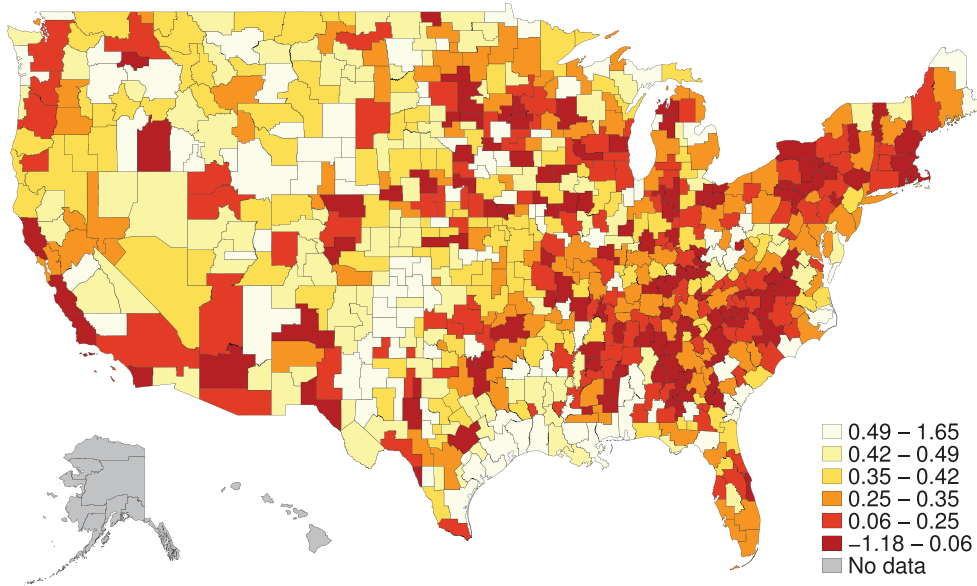


FIGURE 1

Geographical distribution of the welfare effects from the rise of China

Notes: This figure plots the geographic distribution of $100(\hat{W}_g - 1)$, where \hat{W}_g are the welfare effects for group g in the US from the counterfactual rise of China, for our preferred value of $\kappa = 1.5$.

5.3. Import competition and income

In Section 2.6, we showed that changes in relative income can be approximated by our Bartik measure of import competition: $\ln(\hat{Y}_g/\hat{Y}) \approx \frac{1}{\kappa} \ln \sum_s \pi_{gs} \hat{r}_s = -\frac{1}{\kappa} \ln \hat{I}_g$. We check the accuracy of this approximation for the calibrated China shock by comparing the model-implied values for $\ln(\hat{Y}_g/\hat{Y})$ and $\ln \sum_s \pi_{gs} \hat{r}_s$ across groups in the US for the impact of the calibrated China shock. As implied by the approximation, the relationship is almost linear, and the slope is virtually undistinguishable from $1/\kappa$ (see Appendix Figures A.4 and A.5).

This finding implies that $\ln \hat{y}_g \approx \ln \hat{y} + \frac{1}{\kappa} \ln \sum_s \pi_{gs} \hat{r}_s$, which is important for two reasons. First, it confirms that $\frac{1}{\kappa} \ln \sum_s \pi_{gs} \hat{r}_s$ (or $-\frac{1}{\kappa} \ln \hat{I}_g$) can serve as an approximate sufficient statistic for a group's welfare change relative to that for the economy as a whole. This is useful because, in contrast to the exact result in Proposition 1, it does not require knowing the group-level employment changes $\hat{\pi}_{gs}$.⁴⁵

Second, we can test this empirical prediction of the model by regressing changes in CZs' average income on $\ln \sum_s \pi_{gs} \hat{r}_s$, instrumented by the ADH shock.⁴⁶ In line with the model, trade-induced changes in import-competition lead to strong and statistically significant changes in

Figure A.3 shows that dispersion in $\hat{w}_{US,s}$ across sectors converges to zero as κ increases. Finally, notice also that in the final column 2, κ also indirectly affects the multi-sector ACR term, even though $\hat{T}_{China,s}$ is held constant. This is again because κ affects wage changes in all countries and thereby also the changes in expenditure shares $\hat{\lambda}_{jjs}$.

45. As in Kovak (2013), the relationship we find between $\ln \hat{y}_g$ and $\ln \sum_s \pi_{gs} \hat{r}_s$ also provides a theoretical foundation for the empirical use of Bartik-style regressors which assign national sectoral changes to groups based on their initial sectoral composition. Relative to Kovak (2013), our model allows for heterogeneous labour and imperfect mobility across sectors.

46. We employ the same specification we used for our baseline κ estimation (Table 1), except that now the RHS variable is $\ln \sum_s \pi_{gs} \hat{r}_s$ rather than $\ln \hat{\pi}_{NMg}$.

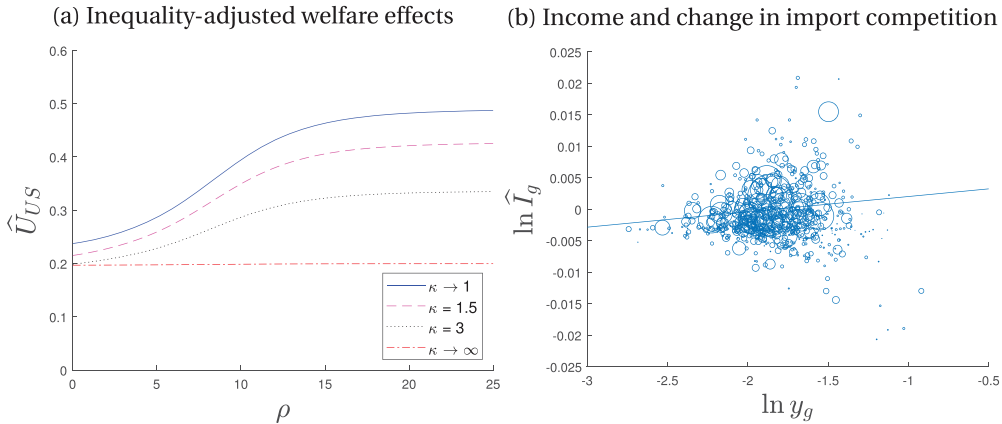


FIGURE 2

Inequality-adjusted welfare effects of the China shock

Notes: (a) The relationship between the inequality-adjusted welfare effects of the rise of China $\hat{U}_{US} \equiv \left(\sum_g \omega_g \hat{W}_g^{1-\rho} \right)^{\frac{1}{1-\rho}}$ and ρ . Here, ρ is the coefficient of relative risk aversion for the agent behind the veil of ignorance and $\omega_g \equiv \frac{l_g (Y_g/L_g)^{1-\rho}}{\sum_h l_h (Y_h/L_h)^{1-\rho}}$ a modified weight for group g . The vertical axis displays $100(\hat{U}_{US} - 1)$. (b) The relationship between $\ln \hat{I}_g = \ln \left(\sum_s \pi_{igs} r_{is} \right)^{-1}$, our measure for the change in regional import-competition (computed for $\kappa = 1.5$), and the logarithm of a group's average income per worker. The solid line displays the linear fit for this relationship, with each commuting zone weighted by its population size. The size of a circle indicates the population size for that commuting zone.

relative income across groups (see [Supplementary Appendix Table D.8](#)). Although high standard errors on the estimated coefficient prohibit us from making strong inferences for the associated values for κ , the implied value for κ is not significantly different from our estimated value of $\kappa = 1.5$.

5.4. Inequality-adjusted welfare effect

We summarize the aggregate and distributional welfare effects of the rise of China for the US by computing the inequality-adjusted welfare effect from equation (19) (see Figure 2a). The consensus in the literature is that plausible values for the coefficient of inequality aversion ρ are between 1 and 3.⁴⁷ For these values and for $\kappa = 1.5$, the inequality-adjusted welfare effect of the rise of China is around 0.23%, which is slightly above the inequality neutral welfare gain of 0.22%. This finding is driven by a positive but small correlation between groups' income and the change in import competition they experience, as is clear from the weighted linear fit between $\ln \hat{y}_g$ and $\ln \hat{I}_g$ in Figure 2b.

6. GAINS FROM TRADE

In this section, we compute the aggregate and group-level gains from trade as described in Section 2, i.e., by computing the negative of the proportional gains from a counterfactual move back to

47. For instance, using agents' intertemporal elasticity of substitution to estimate the curvature parameter, Lucas (2003) argues that $\rho \approx 1$, while a review of the literature leads Hall (2009) to the conclusion that $\rho = 2$. An alternative approach is to calibrate ρ based on people's aversion to risk. Using an indirect approach based on the labour supply elasticity, Chetty (2006) finds that $\rho < 2$, while more direct estimates based on people's decisions under uncertainty range from $\rho = 1$ in Bombardini and Trebbi (2012) to $\rho \approx 3$ in Paravisini, Rappoport and Ravina (2016).

TABLE 3
Aggregate and Group-level Gains from Trade

κ	\widehat{W}_{US}	Mean	CV	Min.	Max.	$\prod_s \widehat{\lambda}_s^{-\beta_s/\theta_s}$
$\rightarrow 1$	1.61	1.61	0.90	-7.95	3.81	1.45
1.5	1.56 (0.05)	1.56 (0.05)	0.63 (0.26)	-4.82 (2.85)	3.03 (0.71)	1.45 (0)
3.0	1.50	1.50	0.33	-1.68	2.24	1.45
$\rightarrow \infty$	1.45	1.45	0	1.45	1.45	1.45

Notes: The first column displays the aggregate gains from trade for the US, in percentage terms ($100(1 - \widehat{W}_{US})$) and the second column shows the mean welfare effect: $100(\frac{1}{G} \sum_g 1 - \widehat{W}_{US,g})$. Here, \widehat{W}_{US} and $\widehat{W}_{US,g}$ are the aggregate and group-level welfare change from a return to autarky for the US. The third column shows the coefficient of variation (CV), and for the fourth and fifth column we have $\text{Min.} = \min_g 100(1 - \widehat{W}_{US,g})$ and $\text{Max.} = \max_g 100(1 - \widehat{W}_{US,g})$, respectively. The final column displays the multi-sector ACR term $100(1 - \prod_s \widehat{\lambda}_{US,US,s}^{-\beta_{US,s}/\theta_s})$. The third row has standard errors in parentheses, computed using the delta method and numerical derivatives with respect to $\hat{\beta} = 1/\hat{\kappa}$, for each statistic when $\kappa = 1.5$. [Supplementary Appendix Table E.3](#) provides a robustness check with sector-specific θ_s values.

autarky. Table 3 summarizes the results. For our estimated value of $\kappa = 1.5$, the aggregate gains from trade with no inequality aversion are 1.56%. As suggested by the theory, the gains from trade decrease with κ , but the effect is small, going from 1.61% for $\kappa = 1$ to 1.45% for $\kappa \rightarrow \infty$.

As in the analysis of the China shock, the main effect of κ is on the distribution of the gains from trade across groups, with the CV decreasing from 0.9 for $\kappa = 1$ to 0 for $\kappa \rightarrow \infty$. For our preferred value of $\kappa = 1.5$, the CV is 0.63, and the range is $[-4.82, 3.03]$. The distribution of gains is skewed to the left with a long tail of low gains, but only 7% of the groups lose from trade (see Appendix Figures A.6 and A.7).

As implied by the analysis above (Sections 2.6 and 5.3), our Bartik measure of import competition $I_g \equiv \sum \pi_{gs} \frac{\beta_s}{r_s}$ perfectly ranks groups in terms of winners and losers from trade for all values of κ (see Appendix Figure A.8). The textile industry faces the highest degree of import competition (with $\frac{\beta_s}{r_s} = 1.52$; Appendix Table A.1), so groups particularly specialized in this industry will gain the least. Interestingly, there is a large region with heavy concentration of groups facing particularly strong import-competition—in part due to specialization in the textile industry—centred around the South-Central and Southern Appalachia regions (see Appendix Figure A.9).

Appendix Figure A.10 shows that for $\rho > 0$, the inequality-adjusted gains from trade are higher than the standard gains, $IGT > GT$, and that IGT increases with ρ . This is a reflection of the fact that, as illustrated in Appendix Figure A.11, most groups at the bottom of the income distribution experience negative degrees of import-competition ($\ln I_g < 0$), due to their specialization in the non-manufacturing sector.⁴⁸

7. EXTENSIONS

7.1. Intermediate goods

Extending the model to allow for an input–output structure is potentially important because a significant share of the value of production in a sector originates from other sectors, and taking this into account may matter for the effects of trade on wages \hat{w}_{is} and welfare across groups. The labour supply of the model is exactly as in the baseline model (see Equation (3)). On the trade side,

48. A regression of $\ln I_g$ on $\ln y_g$, employing population weights, yields a positive and significant coefficient of 0.017, and the R^2 is 14%.

TABLE 4
Counterfactual analysis for the model with intermediates

	\widehat{W}_{US}	Mean	CV	Min.	Max.	ACR
Rise of China	0.37 (0.02)	0.45 (0.04)	0.54 (0.14)	-0.90 (0.35)	1.28 (0.42)	0.29 (0.01)
Gains from trade	2.86 (0.05)	2.92 (0.08)	0.29 (0.12)	-2.12 (2.2)	4.12 (0.62)	2.74 (0)

Notes: The tables show summary statistics for welfare effects of US groups for the model with an input–output structure for $\kappa = 1.5$. The first two rows show results for the counterfactual rise of China, and the final two rows show results for group-level gains from trade. The first column displays the aggregate welfare effect for the US, in percentage terms $100(\widehat{W}_{US} - 1)$ and the second column shows the mean welfare effect: $100(\frac{1}{G} \sum_g \widehat{W}_{US,g} - 1)$. The third column shows the coefficient of variation (CV), and for the fourth and fifth column we have $\text{Min.} = \min_g 100(\widehat{W}_{US,g} - 1)$ and $\text{Max.} = \max_g 100(\widehat{W}_{US,g} - 1)$, respectively. The final column displays the multi-sector ACR term $100\left(\prod_{s,k} \hat{\lambda}_{US,US,k}^{-\beta_{US,s} \tilde{a}_{US,sk} / \theta_s} - 1\right)$. For the gains from trade, we simulate the return to autarky and report the negative of the above statistics for the obtained counterfactual results. Rows 2 and 4 have standard errors, computed using the delta method and the numerical derivatives with respect to $\hat{\beta} = 1/\hat{\kappa}$, in parentheses. [Supplementary Appendix Table F.1](#) has results for other κ values.

the model is identical to [Caliendo and Parro \(2015\)](#), except that wages are now sector-specific (i.e. wages are w_{is} instead of w_i). Hence, trade shares and the price indices are as in equations (1) and (2), but instead of w_{is} we now have c_{is} , where c_{is} is given by $c_{is} = w_{is}^{1-\gamma_{is}} \prod_k P_{ik}^{\gamma_{iks}}$, with

$P_{js} = \zeta_s^{-1} \left(\sum_i T_{is} (\tau_{ijs} c_{is})^{-\theta_s} \right)^{-1/\theta_s}$. The terms γ_{iks} are Cobb-Douglas input shares: a share γ_{iks} of the output of industry s in country i is used buying inputs from industry k , and $1 - \gamma_{is}$ is the share spent on labour, with $\gamma_{is} = \sum_k \gamma_{iks}$. Given this structure, we derive in [Supplementary Appendix F](#) the following expression for a group’s welfare change:

Proposition 3. *Given some trade shock, the percentage change in the real income of group g in country i is given by*

$$\widehat{W}_{ig} = \prod_{s,k} \hat{\lambda}_{iik}^{-\beta_{is} \tilde{a}_{isk} / \theta_s} \cdot \prod_{s,k} \hat{\pi}_{igk}^{-\beta_{is} \tilde{a}_{isk} (1-\gamma_{ik}) / \kappa_{ig}}, \tag{24}$$

where \tilde{a}_{isk} is the typical element of matrix $(I - \Upsilon_i^T)^{-1}$ with $\Upsilon_i \equiv \{\gamma_{iks}\}_{k,s=1,\dots,S}$.

For this extended model, for $\kappa = 1.5$, we find a gain from the China shock of 0.37% and gains from trade of 2.86% (see Table 4).⁴⁹ These gains are higher than in the baseline model, which is in line with the findings in [Costinot and Rodríguez-Clare \(2014\)](#), who explain that the input–output loop in this model leads to an additional round of welfare gains from a given trade shock.

The distributional effects of both the China shock and opening to trade are mitigated compared to the baseline model. The CV is lower in both cases, and the range of group-level welfare effects is slightly more compressed. Still, the correlation between the group-level welfare effects in the two versions of the model is 97.2 for the China shock and 98.3% for the gains from trade (see [Supplementary Appendix Figure F.1](#)).

49. Since the labour supply side of the model is unaltered compared to the baseline model, the κ estimation from Section 4 remains valid. This is why we continue to use the same values for κ in the quantification of this model.

7.2. Imperfect substitutes

In this extension, we introduce two worker types, college and non-college educated workers, so that now there are twice as many groups as in the baseline model (two for each commuting zone). We also allow for the possibility that college and non-college labour are imperfect substitutes, leading to an endogenous college premium that will be affected by trade, similar to the Heckscher–Ohlin model. On the labour supply side, the model remains identical to the baseline model, except that employment shares now have an additional subscript for labour of type $m = C, NC$, for college and non-college workers respectively. On the labour demand side, assuming that efficiency units of college and non-college educated workers enter a CES production function with elasticity of substitution η , then wages satisfy $w_{ims} = \frac{\chi_{ims} Y_{is}}{Z_{ims}}$, where $\chi_{ims} = \frac{\psi_{ims} w_{ims}^{1-\eta}}{\psi_{iCs} w_{iCs}^{1-\eta} + \psi_{iNCs} w_{iNCs}^{1-\eta}}$ is the share of labour type m in total costs in sector s in country i and ψ_{ims} is a corresponding labour-demand shifter.

For the equilibrium analysis, we have market clearing conditions for labour of each type in each industry and country, $ELD_{ims} = 0$, with

$$ELD_{ims} = \chi_{ims} \sum_j \lambda_{ijs} \beta_{js} (Y_j + D_j) - \sum_g w_{ims} Z_{img}.$$

This constitutes a system of equations that we can use to solve for wages, $\{w_{ims}\}$. The following proposition shows the implications for the counterfactual changes in welfare:

Proposition 4. *Given some shock to trade costs or foreign technology levels, the percentage change in the real wage of group mg in country i is given by*

$$\hat{W}_{img} = \prod_s \hat{\lambda}_{iis}^{-\beta_{is}/\theta_s} \cdot \prod_s \hat{\pi}_{img}^{-\beta_{is}/\kappa_{ig}} \cdot \prod_s \hat{\chi}_{ims}^{-\beta_{is}/(\eta-1)}. \quad (25)$$

Compared to Proposition 1, we now have the extra term $\prod_s \hat{\chi}_{ims}^{-\beta_{is}/(\eta-1)}$, which captures the welfare effect of the change in the college premium. If $\kappa_{ig} \rightarrow \infty$ for all ig , then the model collapses to the Heckscher–Ohlin model with gravity as analysed for example in [Burstein and Vogel \(2011\)](#) or [Costinot and Rodríguez-Clare \(2014\)](#). If $\kappa_{ig} \rightarrow 1$ for all ig , then there is no scope for reallocation across sectors within each group-labour type cell, and so all wage changes are at the sector level, $\hat{w}_{iCs} = \hat{w}_{iNCs}$, implying that $\hat{\chi}_{iCs} = \hat{\chi}_{iNCs} = 1$, as in the case considered in the next subsection.

For our quantitative analysis, we set $\eta = 1.6$, as in [Katz and Murphy \(1992\)](#), and similar to [Krusell, Ohanian, Ríos-Rull and Violante \(2000\)](#) and [Acemoglu and Autor \(2011\)](#). How does lowering η from infinity (i.e. perfect substitutes) to this value $\eta = 1.6$ affect the welfare effects of trade? It is instructive to start by focusing on the results under $\kappa \rightarrow \infty$, which implies that there are no changes in the middle, “Roy” term in equation (25), and the only source of heterogeneity is between college and non-college workers due to changes in the college premium as captured by $\prod_s (\hat{\chi}_{iCs} / \hat{\chi}_{iNCs})^{-\beta_{is}/(\eta-1)}$. As shown in Table 5, the rise of China decreases the college premium by 0.03%, while overall trade increases it by 1%. The finding that the college premium falls slightly as a consequence of the rise of China is surprising, but it is explained by a large induced contraction of the electrical and optical equipment industry, which has the second highest cost share of college workers. In contrast, opening up to trade leads to a very large contraction of the textile sector, which has a low cost share of college workers.

As discussed above, moving from $\kappa \rightarrow \infty$ to $\kappa = 1.5$ brings the Roy term to life and softens the effect of trade on the college premium, which now falls by merely 0.01 percent with the rise

TABLE 5
College and non-college workers as imperfect substitutes

(a) The rise of China					
	\widehat{W}_{US}	Mean	CV	Roy gains	College premium
$\kappa \rightarrow \infty; \eta = 1.6$	0.15	0.15	0.15	0.00	-0.03
$\kappa = 1.5; \eta = 1.6$	0.22	0.31	0.77	0.07	-0.01
	(0.02)	(0.04)	(0.17)	(0.03)	(0.02)
$\kappa = 1.5; \eta \rightarrow \infty$	0.22	0.32	0.73	0.07	0.00
(b) Gains from trade					
	\widehat{W}_{US}	Mean	CV	Roy gains	College premium
$\kappa \rightarrow \infty; \eta = 1.6$	1.45	1.48	0.14	0.00	1.00
$\kappa = 1.5; \eta = 1.6$	1.56	1.60	0.47	0.11	0.10
	(0.05)	(0.06)	(0.18)	(0.05)	(0.09)
$\kappa = 1.5; \eta \rightarrow \infty$	1.56	1.49	0.59	0.11	0.00

Notes: The table presents the welfare effects for trade shocks for the model with college and non-college workers as potentially imperfect substitutes. Panel (a) shows results for the rise of China and panel (b) for the gains from trade. The first column provides the parameter values for the simulation results in that row, where $\eta \rightarrow \infty$ implies that all labour is perfectly substitutable. Column 2 display the aggregate welfare effect for all workers in the US, in percentage terms $100(\widehat{W}_{US,m} - 1)$, column 3 shows the mean welfare effect: $100(\frac{1}{G} \sum_g \widehat{W}_{US,mg} - 1)$, and column 4 the coefficient of variation (CV). The fifth column shows the aggregate Roy gains $100(\sum_{mg} (\frac{Y_{img}}{Y_i}) \prod_s \pi_{img}^{-\beta_{is}/\kappa} - 1)$, and the final column the change in the college premium $100(\prod_s (\hat{\chi}_{iCs} / \hat{\chi}_{iNCs})^{-\beta_{is}/(\rho-1)} - 1)$. The China shock is separately calibrated for the parameter values in each row. For the gains from trade, we simulate the return to autarky, and for that simulation report the negative of the aggregate and mean welfare effect. Standard errors for the benchmark results in the second row, computed using the delta method and the numerical derivatives with respect to $\beta = 1/\hat{\kappa}$, in parentheses. [Supplementary Appendix Tables G.2 and G.3](#) have results split by education type when $\eta = 1.6$.

of China, and increases by only 0.1% with overall trade.⁵⁰ In the last row of Table 5, we also show the results with $\kappa = 1.5$ and $\eta \rightarrow \infty$, i.e., where education groups are perfect substitutes. Comparing across perfect and imperfect substitutes, we see very similar results for the effects of the rise of China, and slightly larger mean and lower CV for the gains from trade under imperfect substitutes than perfect substitutes.

7.3. Heterogeneity within commuting zones and trade costs

We can easily introduce more worker types within a commuting zone, not only allowing for heterogeneity in education (as in the previous subsection) but also for differences in age and gender. Here, we revert to the baseline assumption of perfect substitutability in the labour input from different worker types but allow each worker type m to have a potentially different value for κ_m . Moreover, we can also relax the assumption that all goods are costlessly tradable across CZs in the US. This we do by assuming that there are arbitrary trade costs across US states but no trade costs within states, which amounts to treating each US state as if it were a separate country.⁵¹

50. In [Supplementary Appendix Section G](#), we also estimate a separate κ_M for college and non-college workers and examine how this influences the welfare results. Our point estimate for κ_M is somewhat lower for college workers (see [Supplementary Table G.1](#)), and this κ value slightly increases the college premium for both counterfactual scenarios (see [Supplementary Table G.4](#)). In general though, the welfare results are very close to those for the case with a common $\kappa = 1.5$.

51. The quantitative analysis now requires sector-level production and trade data at the level of US states and the other countries, which we borrow from [Rodríguez-Clare, Ulate and Vásquez \(2020\)](#). They build this dataset using data from the Import and Export Merchandise Trade Statistics (from the US Census Bureau), the Commodity Flow Survey

Given the presence of different worker types, we can separately estimate κ_m for each m , employing our baseline regression specification (23). However, due to the within-US trade costs, non-manufacturing wages \hat{w}_{NMt} now vary by US state, which requires the addition of state-by-period fixed effects to our estimation. As discussed in detail in [Supplementary Appendix H](#), the new estimates for κ_m turn out quite similar to those in the baseline.

Although we now have four different groups within each commuting zone, we find that 88% of the variance in the simulated welfare changes across groups is explained by the commuting zone to which they belong. This high share arises from the high correlation in π_{gs} across worker types within a CZ. Hence, the baseline model already captures this driver of the distributional effects fairly well. Still focusing on the distributional effects, we find that the presence of trade costs between US states tends to increase dispersion in the welfare effects compared to the baseline model.⁵² This is mainly due to the ACR welfare term varying across states.

7.4. Mobility across commuting zones

In [Supplementary Appendix J](#), we show how to extend our analysis to allow for mobility of workers across commuting zones. Unfortunately, the data requirements are severe, and we have left this analysis for future work. We note, however, that ADH find insignificant effects of the China shock on population shifts at the commuting zone level, and hence we expect that adding mobility in a way that is consistent with their evidence should not have sizable effects on our results.⁵³

In addition to migration, an alternative form of mobility across regions arises from changes in commuting patterns, as in [Monte, Redding and Rossi-Hansberg \(2018\)](#). Extending our model to allow for commuting and exploring the impact of the China shock in that setting is an interesting task but beyond the scope of this article. Here, we simply note that, as shown in [Section 5](#), the regions that are most negatively affected by the China shock tend to be geographically concentrated, and so commuting is unlikely to serve as a significant margin of adjustment.

8. EMPLOYMENT EFFECTS

In this section, we extend the model so that total employment is endogenous both because of the possibility of home production, modelled as in [Caliendo *et al.* \(2019\)](#), and because of involuntary unemployment due to search and matching frictions, modelled as in [Kim and Vogel \(2021\)](#). We then estimate the model and report the group-level and aggregate effects of the China shock.

(CFS), and the Regional Economic Accounts of BEA Commodity Flow Service. In principle, the model could allow for trade costs between geographical units at an even more disaggregated level, but we are not aware of reliable data at lower levels of disaggregation that cover the entire US.

52. This finding comes out clearest in a version of the model where we only add trade costs and don't consider different groups within each CZ, such that we have a clean comparison with the results of the baseline model. The results for this version of the model are discussed in detail in the 2020 version of this article.

53. [Caliendo *et al.* \(2019\)](#) and [Adão, Arkolakis and Esposito \(2020\)](#) allow for mobility across both sectors and regions and quantify the effect of the China shock at the level of US states rather than commuting zones (CZs). Their results also point to weak effects of trade shocks on mobility across regions. Relatedly, the reduced form evidence on the regional migration response to trade shocks in the US is mixed. For instance, [Greenland, Lopresti and McHenry \(2019\)](#) find a substantial impact on CZs' population growth arising from the granting of Permanent Normal Trade Relations to China, primarily driven by adjustments among the younger cohorts. In contrast, [Choi, Kuziemko, Washington and Wright \(2020\)](#) find no local US migration response in the US after the introduction of NAFTA.

8.1. Model

There are three periods. In the first one, workers learn about their productivity in home production and formal employment and decide whether to seek formal employment based on the expected income in each of those options. In the second period, workers who chose formal employment learn about their sector-specific productivity realization and decide in which sector to apply for work. This decision depends on the probability of employment and the wage per efficiency unit in each sector. In the third period, workers learn whether they are employed or unemployed.⁵⁴

For each worker, productivity in home production is z_{HP} while productivity in sector s is $z_F z_s$. As in Section 2.1, the productivity terms z_s are drawn independently from a Fréchet distribution with shape and scale parameters κ and A_{igs} , respectively. In addition, z_{HP} and z_F are drawn independently from a Fréchet distribution with shape and scale parameters μ , A_{igHP} , and A_{igF} , respectively. Finally, firms can post vacancies at a cost of c in terms of the final good (i.e. aggregating across sectors in the same way that consumers do) and capture an exogenous share $1 - \nu$ of the value generated from a match. Here, total sector-level matches (M_{igs}) are a Cobb–Douglas function of vacancies (V_{igs}) and labour supply (L_{igs}): $M_{igs} = A_{igM} V_{igs}^\alpha L_{igs}^{1-\alpha}$, with $\alpha \in (0, 1)$.⁵⁵

With free entry of firms to posting vacancies in each sector, in equilibrium, we must satisfy the zero-profit condition

$$cV_{igs} = (1 - \nu)\omega_{is}e_{igs}Z_{igs} \tag{26}$$

for all s , where e_{igs} is the employment rate ($e_{igs} \equiv M_{igs}/L_{igs}$) in sector s , ω_{is} the real wage in sector s , and Z_{igs} the total efficiency units of labour supplied to sector s .⁵⁶ The results of Section 2.1 still apply so that we have $\omega_{is}Z_{igs}/L_{igs} = \xi \Phi_{ig}\bar{z}_{igF}$ and $L_{igs} = \pi_{igs}L_{igF}$, with π_{igs} , Φ_{ig} and ξ as in Section 2.1 (except with $\nu\omega_{is}e_{igs}$ now playing the role of w_{is}) and where L_{igF} is the number of workers seeking formal employment and $\bar{z}_{igF} \equiv \frac{Z_{igF}}{L_{igF}}$ is the corresponding average efficiency units per worker. Combining the expressions for the matching function, the zero-profit condition, the employment rate and revenue per applicant, we get that the employment rate is common across sectors, $e_{igs} = e_{ig}$ for all s , and is given by

$$e_{ig} = A_{igM}^{\frac{1}{1-\alpha}} \left(\frac{(1-\nu)\xi}{c} \right)^{\frac{\alpha}{1-\alpha}} (\Phi_{ig}\bar{z}_{igF})^{\frac{\alpha}{1-\alpha}}. \tag{27}$$

The result that the employment rate is common across sectors depends of course on our assumption of no cross-sector variation in ν , α , and c within a group.

The only remaining task is to solve for L_{igF} and \bar{z}_{igF} . Since workers make these decisions based on the expected value of formal employment, $\eta\nu e_{ig}\Phi_{ig}$, then we know from the standard

54. An alternative approach is to assume a two-period structure, with a nested Fréchet distribution for productivity draws in home production and in each of the formal sectors. The problem with this specification is that it would require the elasticity of substitution between home production and formal employment to be lower than the one across formal sectors, which is not what the data implies.

55. Our theoretical results remain valid if the parameters α , μ , κ , ν , and c vary across groups. Thus, as in Section 2, we could allow these parameters to vary across groups and write them with the ig subscript. However, here we choose not to do that to ease the notational burden. In any case, when we come to estimation, we will need to assume that α , μ , and κ are common across groups in the US.

56. We have assumed here that unemployment goes along with no income, so that the surplus of a match is just $\omega_{is}z_s z_F$. We could instead assume that there are unemployment benefits financed from a common tax rate on employed workers. Since being employed is randomly determined, and assuming that the tax rate is common across sectors, these benefits have no distortive effects, and all our results remain valid.

Fréchet algebra that $L_{igF} = \pi_{igF} L_{ig}$, with

$$\pi_{igF} = \frac{A_{igF} (\xi v e_{ig} \Phi_{ig})^\mu}{A_{igHP} \omega_{iHP}^\mu + A_{igF} (\xi v e_{ig} \Phi_{ig})^\mu}, \quad (28)$$

where ω_{iHP} is the exogenous real wage per efficiency unit in home production. Expected welfare (and average real income among all workers) is

$$W_{ig} = \tilde{\xi} (A_{igHP} \omega_{iHP}^\mu + A_{igF} (\xi v e_{ig} \Phi_{ig})^\mu)^{1/\mu}, \quad (29)$$

where $\tilde{\xi} \equiv \Gamma(1 - 1/\mu)$. By the properties of the Fréchet distribution, this is also the average real income among all workers choosing formal employment. Applying the same logic as in Section 2.1, we can show that $W_{ig} = \xi v e_{ig} \Phi_{ig} Z_{igF} / L_{igF}$, and hence $\bar{z}_{igF} \equiv \frac{Z_{igF}}{L_{igF}} = \frac{W_{ig}}{\xi v e_{ig} \Phi_{ig}}$. This implies that

$$e_{ig} = A_{igM} \left(\frac{(1-v)}{vc} \right)^\alpha W_{ig}^\alpha. \quad (30)$$

We assume that parameters are such that the solution to this equation entails $e_{ig} \in (0, 1)$. The equilibrium system to solve for all wages is very similar to the one for the baseline model (see [Supplementary Appendix I.1](#)).

Proposition 5. *Given some trade shock, the percentage change in the real income of group g in country i is given by*

$$\hat{W}_{ig} = \left(\pi_{igHP} + (1 - \pi_{igHP}) \hat{e}_{ig}^\mu \hat{\Phi}_{ig}^\mu \right)^{(1/\mu)}, \quad (31)$$

where $\hat{\Phi}_{ig}$ captures country and group-level gains from specialization

$$\hat{\Phi}_{ig} = \prod_{s \in F} \hat{\lambda}_{iis}^{-\frac{\beta_{is}}{\theta_s}} \prod_{s \in F} \hat{\pi}_{igs}^{-\frac{\beta_{is}}{\kappa}},$$

and where the change in the employment rate comes from the solution to

$$\hat{e}_{ig}^{\mu/\alpha} = \pi_{igHP} + (1 - \pi_{igHP}) \hat{e}_{ig}^\mu \hat{\Phi}_{ig}^\mu. \quad (32)$$

It is easy to verify that a trade shock leads to a change in employment in the same direction as in Φ_{ig} , so that $\hat{e}_{ig} < 1$ if $\hat{\Phi}_{ig} < 1$ and $\hat{e}_{ig} > 1$ if $\hat{\Phi}_{ig} > 1$.

Intuitively, a negative trade shock leads to a decline in the real wage, which makes posting vacancies less profitable because the cost is in terms of the final good and the benefit is in terms of the nominal wage. Via the zero-profit condition, this leads to fewer posted vacancies and a higher unemployment rate, amplifying the effect of trade shocks on welfare. We can see this most clearly if we ignore home production by setting $\pi_{igHP} = 0$. In that case, we would have $\hat{e}_{ig} = \hat{\Phi}_{ig}^{\frac{\alpha}{1-\alpha}}$ and hence $\hat{W}_{ig} = \hat{\Phi}_{ig}^{\frac{1}{1-\alpha}}$, implying an amplification of trade shocks on welfare by the factor $\frac{1}{1-\alpha} > 1$.⁵⁷

57. Amplification through endogenous unemployment arises in a similar way as with an input-output loop: whereas the factor of amplification there is the inverse of the labour share in the production of final goods, here it is the inverse of the labour share in the production of matches, i.e., $1/(1-\alpha)$. In fact, if we had a single sector then the model above would be isomorphic to the [Eaton and Kortum \(2002\)](#) model with an input-output loop where final output is used together with labour to produce final goods according to a Cobb–Douglas production function with labour share $1-\alpha$.

In contrast, home production softens the effect of trade shocks on welfare. This happens first because workers have the option to engage in home production, where the real wage is not affected by the trade shock, and second because the decline in labour supply reduces the effect of the trade shock on the unemployment rate.⁵⁸ Thus, as emphasized by [Kim and Vogel \(2020\)](#), although both home production and frictional unemployment imply that a negative trade shock lowers employment, they have opposite effects on the welfare effects of a trade shock: home production serves as an efficient adjustment mechanism that mitigates the effect whereas frictional unemployment amplifies it.⁵⁹

8.2. Estimation

Dropping the country subscript, we start from the fact that if W_g is real average income among workers in the labour force, then average nominal income among employed workers is $y_g = W_g P / e_g$. Combining this with equations (29) and (30), we obtain

$$\ln \hat{y}_g = \ln \hat{P} + \frac{1-\alpha}{\alpha} \ln \hat{e}_g - \ln \left(\hat{A}_g^M \right)^\alpha, \tag{33}$$

where without loss of generality we have assumed that $\hat{c} = \hat{v} = 1$. Next, combining $y_g = W_g P / e_g$ with equations (29) and (28), and using $\pi_{gHP} = 1 - \pi_{gF}$ yields

$$\ln \left(\hat{e}_g \hat{y}_g \right) = \ln \left(\hat{P} \hat{\omega}_{HP} \right) - \frac{1}{\mu} \ln \hat{\pi}_{gHP} + \ln \hat{A}_{gHP}. \tag{34}$$

Finally, combining $y_g = W_g P / e_g$ with equations (29) and (28), and proceeding as in Section 4 to use $\pi_{gNM} = \frac{A_{gNM} \omega_{NM}^\kappa}{\Phi_g^\kappa}$, we get

$$\ln \left(\hat{y}_g \hat{\pi}_{gF}^{1/\mu} \right) = \ln \left(\hat{P} \hat{\omega}_{NM} \right) - \frac{1}{\kappa} \ln \hat{\pi}_{gNM} + \ln \left(\hat{A}_{gF}^{1/\mu} \hat{A}_{gNM}^{1/\kappa} \right). \tag{35}$$

Equation (35) is analogous to equation (21), with the difference that now the dependent variable is the log of $\hat{y}_g \hat{\pi}_{gF}^{1/\mu}$ rather than the log of \hat{y}_g . The reason for the difference is that now we need to take into account that workers can mitigate the effect of a shock through changes in labour force participation as captured by $\hat{\pi}_{gF}^{1/\mu}$. Equation (34) is also quite similar: all else equal, a higher $\hat{\pi}_{gHP}$ leads to a lower expected income ($\hat{e}_g \hat{y}_g$) with elasticity $1/\mu$. Finally, equation (33) comes from the fact that an increase in the real income of employed workers goes along with an increase in the employment rate with elasticity $\alpha/(1-\alpha)$.

58. To understand this second effect, we can log-linearize equation (32) around $\hat{\Phi}_{ig} = 1$, which implies

$$\frac{d \ln \hat{e}_{ig}}{d \ln \hat{\Phi}_{ig}} = \frac{\alpha}{1-\alpha} \left(\frac{1 - \pi_{igHP}}{1 + \frac{\alpha}{1-\alpha} \pi_{igHP}} \right).$$

This implies that $\frac{d \ln \hat{e}_{ig}}{d \ln \hat{\Phi}_{ig}} |_{\pi_{igHP} > 0} < \frac{d \ln \hat{e}_{ig}}{d \ln \hat{\Phi}_{ig}} |_{\pi_{igHP} = 0}$.

59. One could imagine that frictional unemployment matters for the transmission of trade shocks to welfare because of the inefficiency in vacancy posting whenever the Hosios condition (i.e. $1 - \nu = \alpha$, see [Hosios, 1990](#)) is not satisfied. This is not the case, however, as revealed by the fact that the amplification above is not dependent on the difference between $1 - \nu$ and α . The reason that this inefficiency is irrelevant for the comparative statics of welfare is that the shock does not affect the share of final output that is used for vacancy posting, which is fixed at $1 - \nu$ given the zero-profit condition.

TABLE 6
GMM Estimation of the Model with Unemployment and Labour Force Participation

	(1)	(2)	(3)	(4)
$\ln \hat{E}$	0.985* (0.563)	2.398* (1.456)	4.389 (3.199)	1.002** (0.408)
$\ln \hat{\pi}_{HP}$	-0.312** (0.122)	-0.519 (0.369)	-0.760 (0.708)	-0.362*** (0.104)
$\ln \hat{\pi}_{NM}$	-0.633** (0.291)	-0.972** (0.409)	-1.001** (0.409)	-0.834** (0.258)
Implied α	0.504	0.294	0.186	0.499
Implied μ	3.209	1.927	1.316	2.759
Implied κ	1.579	1.029	0.999	1.199
Observations	1,444	1,444	1,444	1,444
Import penetration	Other (lagged)	Other (no lag)	US (no lag)	US Bartik

Notes: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors are clustered at the state level, and we weight by 1990 commuting zone populations. All log changes for the 2000–7 period are multiplied by (10/7) to obtain decade-equivalent changes. Home production defined as not in the labour force (NiLF).

We use equations (33)–(35) to estimate α , μ , and κ using a standard GMM approach, exploiting the cross-equation restriction on $1/\mu$ in equations (34) and (35).⁶⁰ We employ our standard instrumental variables Z_g (different types of CZ-group-level China shocks or the Bartik instrument). As a natural extension of the identifying assumption in our baseline estimation, we assume that Z_g is uncorrelated to the vector of error terms ϵ_g (i.e. $\mathbb{E}(Z_g' \epsilon_g) = 0$). Intuitively, this means that our instruments are uncorrelated with any group-level shocks that could affect earnings or employment (i.e. unobserved productivity and labour supply/demand shocks).

We estimate this model using an extended version of our baseline data, which includes group-level employment and labour force participation rates. For this exercise, individuals are classified under home production when they are not in the labour force. The results from our estimation are shown in Table 6, with each column representing a separate estimation based on our four instruments. Our estimates for κ are slightly lower than those from the baseline model, while those for α range between 0.2 and 0.5, which is consistent with estimates reviewed in Petrongolo and Pissarides (2001). In the following subsection, we show how these new estimates translate into aggregate and distributional welfare effects.

8.3. Quantitative implications

Based on the estimation results in the previous subsection, we now explore the quantitative implications of the model with $\kappa = 1.5$ and home production and unemployment with $\alpha = 1/3$ and $\mu = 2.5$. We proceed as in the previous sections by calibrating the China shock and then using the model to quantify its effects across commuting zones in the US.⁶¹

60. Specifically, we estimate the following system of equations

$$\begin{pmatrix} \ln \hat{y}_g \\ \ln(\hat{e}_g \hat{y}_g) \\ \ln \hat{y}_g \end{pmatrix} = \underbrace{\begin{pmatrix} x_g & 0 & 0 \\ 0 & x_g & 0 \\ 0 & 0 & x_g \end{pmatrix}}_{\equiv X_g} \Gamma + \underbrace{\begin{pmatrix} \ln \hat{e}_g & 0 & 0 \\ 0 & \ln \hat{\pi}_{gHP} & 0 \\ 0 & \ln \hat{\pi}_{gF} & \ln \hat{\pi}_{gNM} \end{pmatrix}}_{\equiv W_g} \beta + \underbrace{\begin{pmatrix} \epsilon_g^M \\ \epsilon_g^{HP} \\ \epsilon_g^{NM} \end{pmatrix}}_{\equiv \epsilon_g},$$

where $\beta = (\frac{1-\alpha}{\alpha}, -1/\mu, -1/\kappa)'$, X_g is the set of ADH controls from our baseline estimation (including intercepts that are not constrained by any cross-equation restriction), and the vector ϵ_g of error terms (which differ from $(\ln(\hat{A}_g^M)^\alpha, \ln \hat{A}_{gHP}, \ln(\hat{A}_{gF}^{1/\mu} \hat{A}_{gNM}^{1/\kappa}))$ due to the inclusion on the ADH controls).

61. We use the employment data for other countries from (The World Bank, n.d.).

TABLE 7
Welfare effects with and without frictional unemployment and home production

	The rise of China			Gains from trade		
	\widehat{W}_{US}	Mean	CV	\widehat{W}_{US}	Mean	CV
No SAM, no HP	0.2150	0.294	1.010	1.556	1.556	0.626
With SAM, no HP	0.3240	0.443	1.001	2.325	2.327	0.616
With SAM & HP, $\mu = 1$	0.2116	0.288	1.022	1.531	1.529	0.626
With SAM & HP, $\mu = 2.5$	0.2121	0.288	1.011	1.521	1.520	0.618

Notes: The first three columns display the welfare effects for the counterfactual rise of China, while the final three columns show the gains from trade. Columns 1 and 4 displays, for the relevant worker type, the aggregate welfare effect in percentage terms $100(\widehat{W}_{US} - 1)$, and columns 2 and 5 show the mean welfare effect: $100(\frac{1}{G} \sum_g \widehat{W}_{US,g} - 1)$. The third column shows the coefficient of variation (CV). For the gains from trade, we simulate the return to autarky, and for that simulation report the negative of the above statistics. Full results are available in [Supplementary Appendix Tables I.5 and I.6](#). We compute standard errors for these counterfactual exercises in [Supplementary Appendix Table I.7](#).

As expected from the theory discussion and the estimation results, commuting zones more exposed to the China shock experience declines in employment both due to an increase in unemployment and a decline in labour force participation (see [Supplementary Appendix Table I.5](#)). Interestingly however, overall employment increases because the China shock leads to an increase in the average real wage across US commuting zones. This again shows how the reduced-form results in ADH are indicative of relative effects across commuting zones differentially exposed to the China shock, but cannot tell us about the average effect, which here is generated via simulation from the calibrated general equilibrium model.

Turning to welfare, there are three ways in which search and matching and home production affect the welfare effects of trade shocks. First, search and matching leads to amplification, roughly by increasing welfare effects by 50%. Second, once we introduce home production, we change the notion of welfare, where the part of welfare that comes from home production is not affected by trade, which dampens welfare changes roughly by the share of employment in home production. Finally, allowing for movement in and out of home production leads to more favourable welfare effects since people can go into home production if real wages decrease (dampening the effect of the shock) and come out of home production if real wages increase (amplifying the positive effect of the shock).⁶²

We see these effects play out as expected in [Table 7](#). The first row shows the baseline, the second row adds endogenous unemployment via search and matching, the third row adds home production but with $\mu = 1$, and finally the last row shows the results for the full model with the calibrated value of μ . As we move from $\mu = 1$ to $\mu = 2.5$, the effect of the China shock becomes slightly more positive and there is less dispersion in the welfare effects. We also see that the gains from trade slightly fall with μ , which is because the effect of the return to autarky is less negative. Overall however, the welfare results in the full model with endogenous unemployment and labour-force participation are not significantly different from those in the baseline model.

This version of the model generates predictions for the impact of the China shock on income, unemployment, and employment sectoral shares. As an analysis of model fit, we regress actual changes in the data for the period 2000–7 on simulated changes for the employment rate, \hat{e}_g and the share of employment in home production, $\hat{\pi}_{gHP}$. We find that, for both these variables, there is a positive relationship between the simulated and the actual changes (see [Appendix Table A.2](#), columns 1 and 2), although the coefficient on the employment rate is imprecisely estimated. The

62. There may be a negative feedback effect here since a more elastic labour supply may also soften the effect of the shock on equilibrium real wages, but one would expect that such negative feedback effects would not overturn the mechanism described here for how $\mu > 0$ affects welfare in the counterfactual analysis.

model performs slightly better if we increase the employment rate elasticity α from 1/3 to 1/2, which is consistent with our estimation results in Table 6.⁶³ In that case, both point estimates are significantly different from zero and insignificantly different from unity.^{64,65}

9. CONCLUSION

We think of this article as establishing a bridge between two separate literatures. On the one hand, a recent wave of empirical work exemplified most prominently by Autor *et al.* (2013) has shown that trade shocks have important distributional implications, but without deriving welfare effects. On the other hand, research surveyed in Costinot and Rodríguez-Clare (2014) shows how to quantify the welfare effects of trade for a wide class of gravity models, but with so far little to say about distributional implications.⁶⁶ In this paper we extend the multi-sector gravity model of trade to allow for heterogeneous labour as in Roy (1951) and Lagakos and Waugh (2013) and with multiple groups of *ex ante* identical workers as in Burstein *et al.* (2019), and use the resulting framework to derive a simple approach to computing group-level and aggregate welfare effects of trade shocks. We borrow the identification strategy proposed by Autor *et al.* (2013), but we use it here to estimate the model's key parameter governing the degree of labour heterogeneity and the distributional implications of trade shocks.

We use the model to quantify the welfare effects of the China shocks on groups in the US defined as commuting zones. We find that the average effect is positive, that some groups experience losses roughly equal to four times as high as the average gain, and that those groups tend to be concentrated in the Midwest and the inland Eastern region of the US. At the same time, the burden of adjustment to the China shock is spread relatively equally across poor and rich groups. As a consequence, adjusting the welfare calculation for plausible levels of inequality aversion leads to only mild deviations from the standard aggregate effect. Extending our baseline model to allow for intermediate goods, within-country trade costs, heterogeneity within commuting zones, and endogenous employment does not substantially change these conclusions.

The question addressed in this article is complex, and our approach has obvious limitations. Most importantly, our analysis is silent on the effect of shocks on individuals within each group. We deal with this partially by considering finer groups—for example differentiating within a commuting zone by gender, age and education—but even then our approach fails to take into account the large costs of trade-induced layoffs to individual households in the absence of a proper safety net (see e.g. Autor *et al.*, 2014; Pierce and Schott, 2020). Our approach also leaves out the role of nominal wage frictions and endogenous trade imbalances, both of which can affect the path of unemployment after a trade shock as shown in Rodríguez-Clare *et al.* (2020) and Dix-Carneiro, Pessoa, Reyes-Heroles and Traiberman (2021). Understanding the relative importance of all these features in affecting the aggregate and distributional effects of trade shocks is an important challenge for future research.

63. In the baseline we conservatively set $\alpha = 1/3$ to stay closely in line with the values in the literature.

64. Our model predictions also qualitatively match the patterns in the data for changes in income and the employment share in manufacturing. However, our one factor model does not account for changes in the labour share of revenue, which matters for the quantitative fit (Galle and Lorentzen, 2021).

65. As a sanity check, we also regress the model's predicted changes in employment and income per worker on ADH's China shock IV (see Appendix Table A.3). As expected, the China shock is positively associated with the predicted changes in home production shares, and negatively with changes in the employment rate and income per worker. All these correlations are strongly significant.

66. The only mention of distributional implications in Costinot and Rodríguez-Clare (2014) is in regards to Burstein and Vogel (2017), which is limited to quantifying welfare effects among low- and high-skilled workers.

Appendix A. Supplementary Tables and Figures for the Counterfactuals

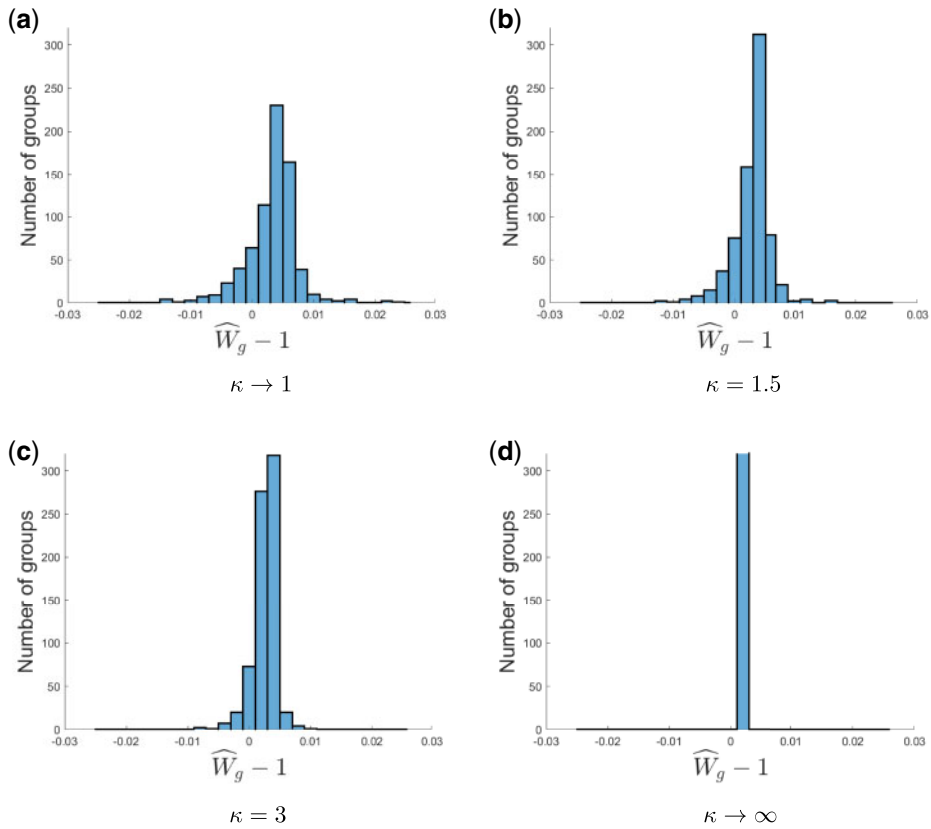


FIGURE A.1

Distribution of the welfare effects for the rise of China

Notes: This figure plots the distribution of $\hat{W}_g - 1$, where \hat{W}_g are the welfare effects for all US groups from the counterfactual rise of China. The different panels show the welfare results for different values of κ , indicated at the bottom of each panel. The vertical axis counts the number of groups in each bin, and the total number of groups is 1,444. For visual reasons, the scale of the vertical axis is censored at 300.

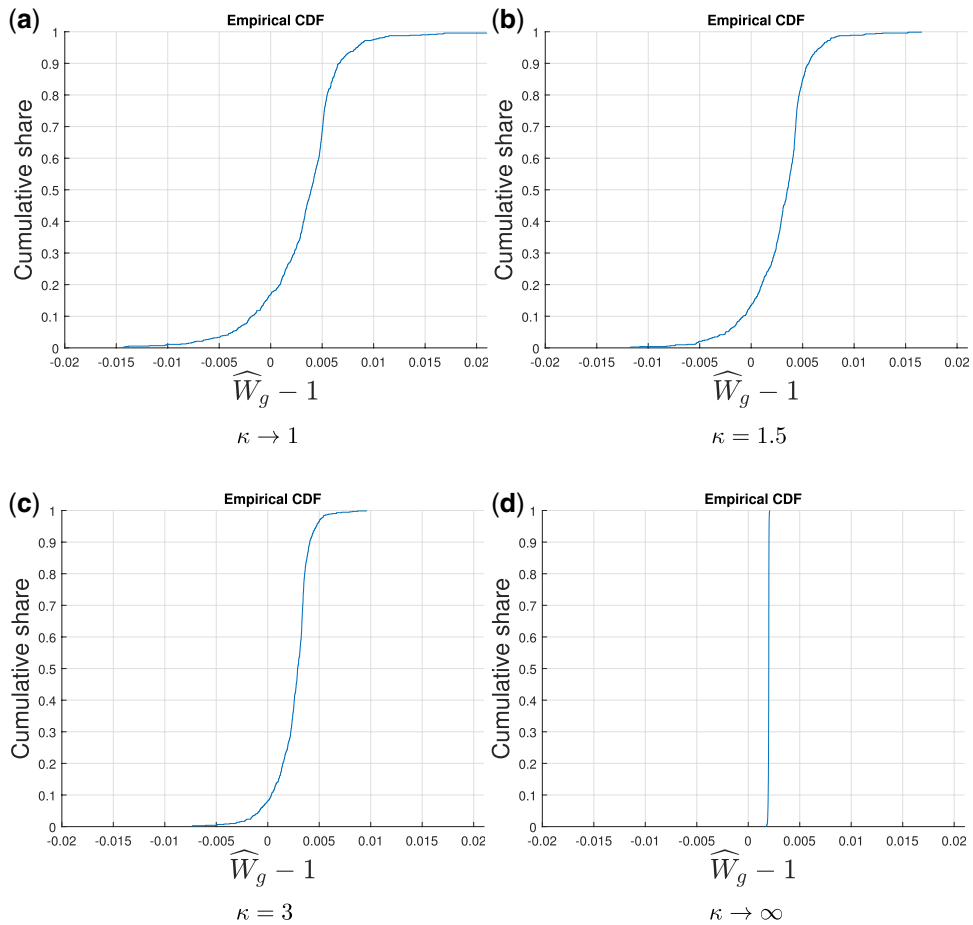


FIGURE A.2

Distribution of the welfare effects for the rise of China

Notes: This figure plots the cumulative density function of $\widehat{W}_g - 1$, where \widehat{W}_g are the welfare effects for all US groups from the counterfactual rise of China. The different panels show the welfare results for different values of κ , indicated at the bottom of each panel.

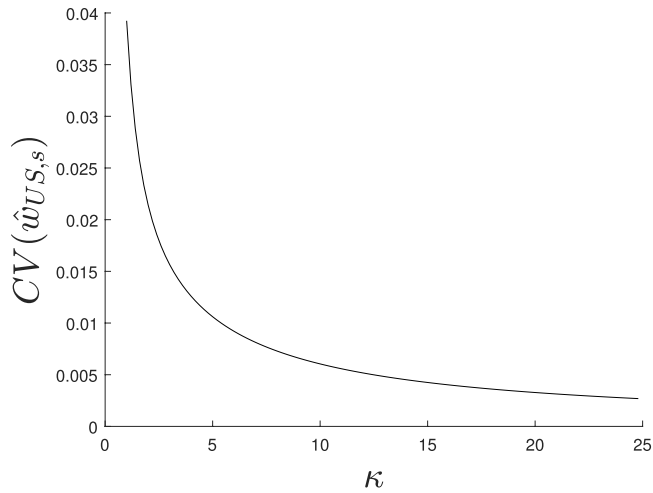


FIGURE A.3

Equilibrium impact of κ on wage changes

Notes: The figure plots the coefficient of variation for wage changes in the US ($\hat{w}_{US,s}$) for a given China shock, as a function of κ .

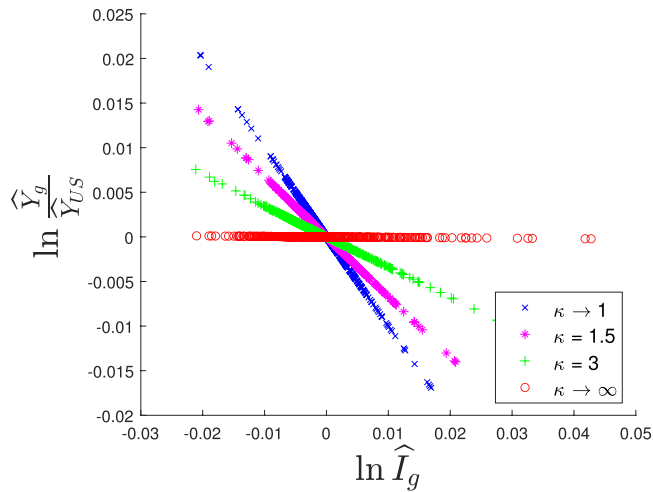


FIGURE A.4

Changes in import competition and groups' relative income for the China shock

Notes: The figure plots the value for $\ln \frac{\hat{Y}_g}{\hat{Y}_{US}}$ in relation to $\ln \hat{I}_g = -\ln \sum_s \pi_{gs} \hat{r}_s$, our Bartik measure for the change in groups' import-competition. Each scatter represents the simulation results for a different value of κ , for values of $\hat{T}_{China,s}$ calibrated for $\kappa = 1.5$.

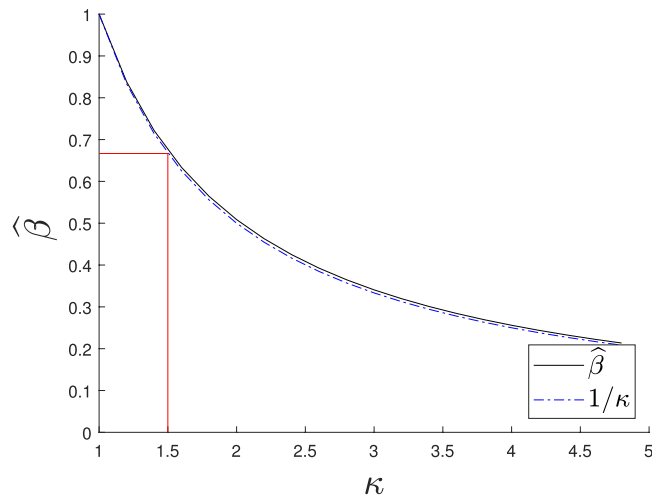


FIGURE A.5

A Bartik approximation of income changes

Notes: The coefficient $\hat{\beta}$, on the vertical axis, is estimated in the following regression: $\ln \hat{y}_g = \alpha + \beta \ln \sum_s \pi_{gs} \hat{r}_s + \varepsilon_g$, which is run separately for different sets of simulation outcomes for \hat{y}_g and \hat{r}_s . Each set of simulation outcomes is obtained for a different value of κ (horizontal axis). The vertical line represents the preferred value for κ from the structural estimation in Section 4, and the solid horizontal line represents the associated value for β . Also note that Appendix Figure A.4 shows that the relation between the model-implied values for $\ln(\hat{Y}_g/\hat{Y})$ and $\ln \sum_s \pi_{gs} \hat{r}_s$ across groups in the US for the impact of the calibrated China shock is (almost) exactly linear.

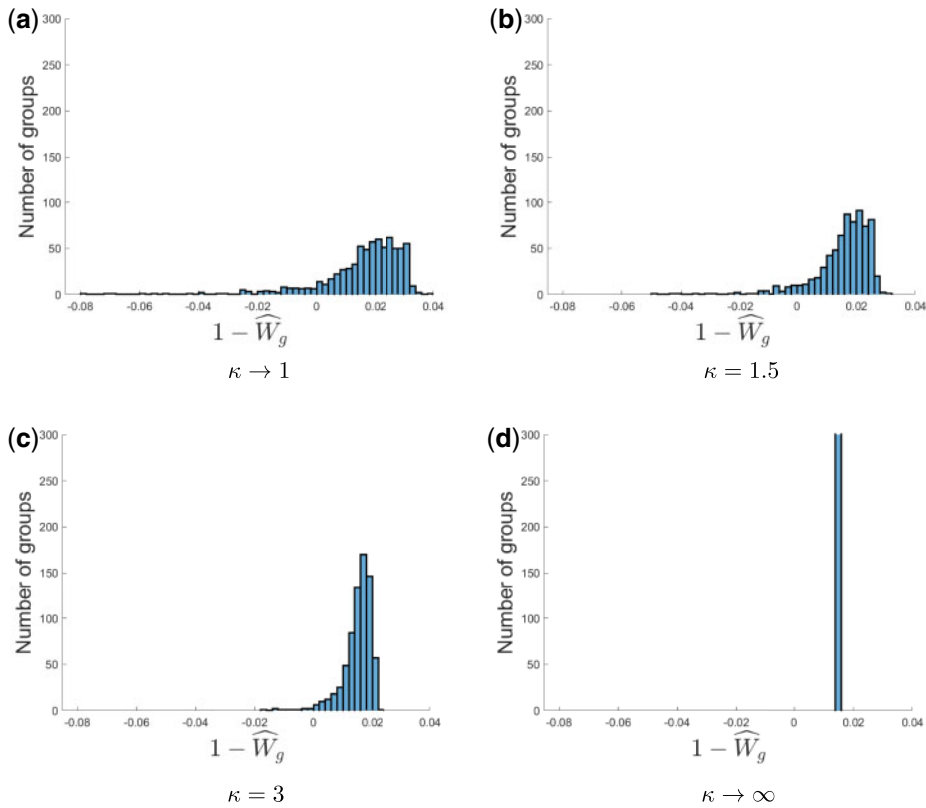


FIGURE A.6

Distribution of the gains from trade

Notes: This figure plots the distribution of $1 - \widehat{W}_g$, where \widehat{W}_g are the welfare effects for all US groups from a return to autarky. The different panels show the welfare results for different values of κ , indicated at the bottom of each panel. The vertical axis counts the number of groups in each bin, and the total number of groups is 1,444. For visual reasons, the scale of the vertical axis is censored at 300.

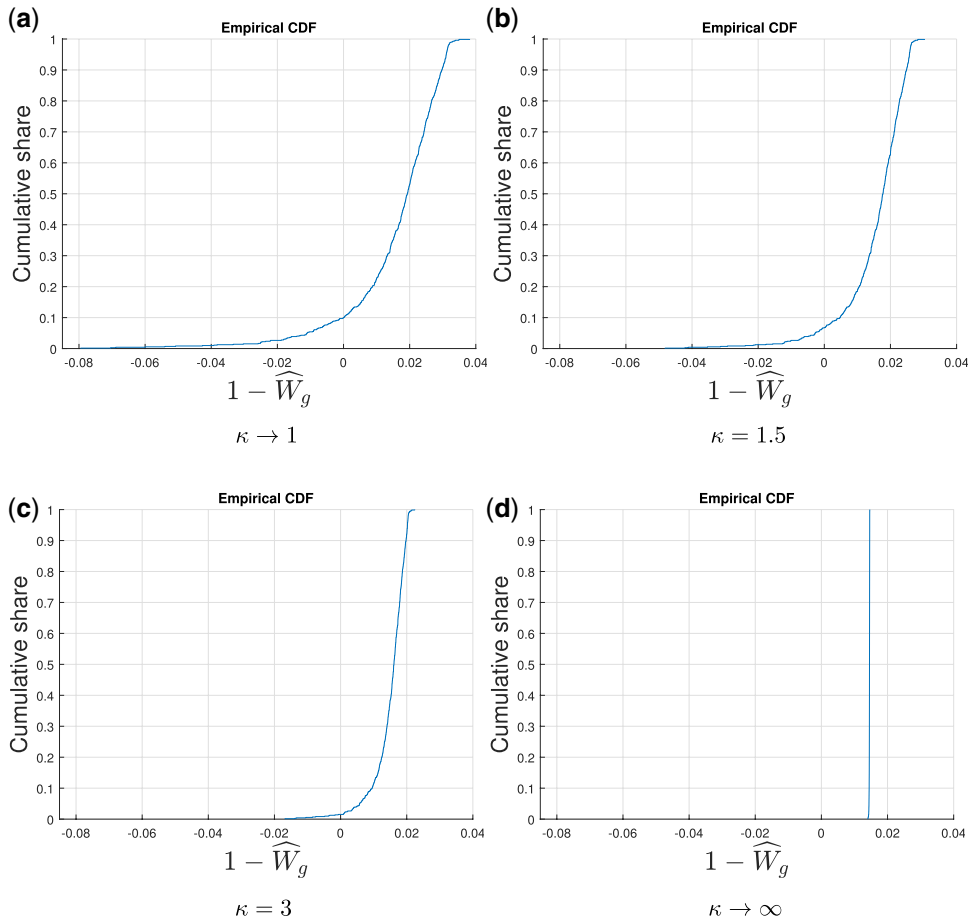


FIGURE A.7

Cumulative density functions for the gains from trade

Notes: This figure plots the cumulative density function of $1 - \widehat{W}_g$, where \widehat{W}_g are the welfare effects for all US groups from a return to autarky. The different panels show the welfare results for different values of κ , indicated at the bottom of each panel. (a) $\kappa \rightarrow 1$, (b) $\kappa = 1.5$, (c) $\kappa = 3$, and $\kappa \rightarrow \infty$.

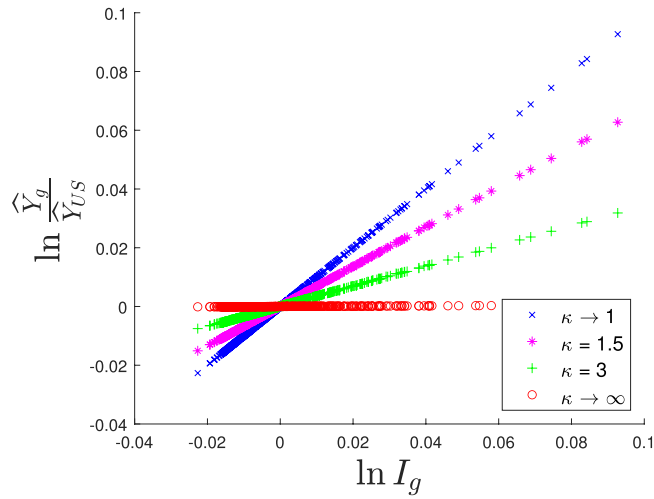


FIGURE A.8

Import competition and groups' relative gains from return to autarky

Notes: The figure plots the value for $\ln \frac{\hat{Y}_g}{Y_{US}}$ in relation to $\ln I_g = \sum_s \pi_{igs} \frac{\beta_{is}}{r_{is}}$, our Bartik measure for groups' import-competition. Each scatter represents the simulation results for the return to autarky for a different value of κ .

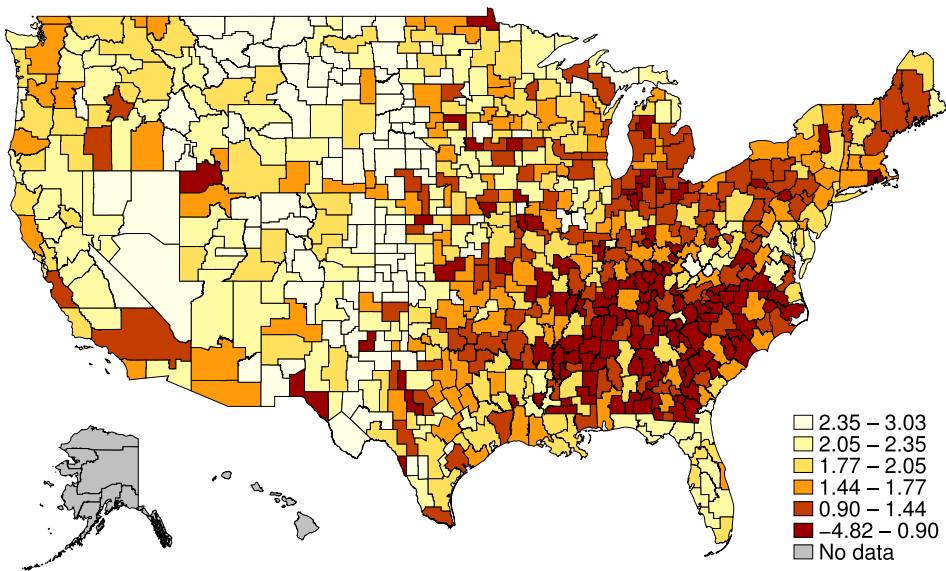


FIGURE A.9

Geographical distribution of the gains from trade

Notes: This figure plots the geographic distribution of $100(1 - \hat{W}_g)$, where \hat{W}_g are the welfare effects for group g in the US from a return to autarky for our preferred value of $\kappa = 1.5$.

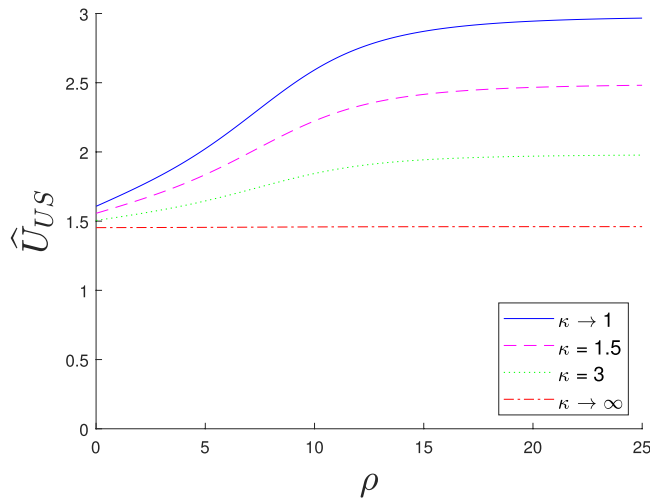


FIGURE A.10

Inequality-adjusted gains from trade

Notes: The figure plots the relationship between the inequality-adjusted gains from trade $\hat{U}_{US} \equiv \left(\sum_g \omega_g \hat{W}_g^{1-\rho} \right)^{\frac{1}{1-\rho}}$ and ρ . Here, ρ is the coefficient of relative risk aversion for the agent behind the veil of ignorance and $\omega_g \equiv \frac{l_g (y_g/L_g)^{1-\rho}}{\sum_h l_h (y_h/L_h)^{1-\rho}}$ a modified weight for group g . The vertical axis displays $100(1 - \hat{U}_{US})$.

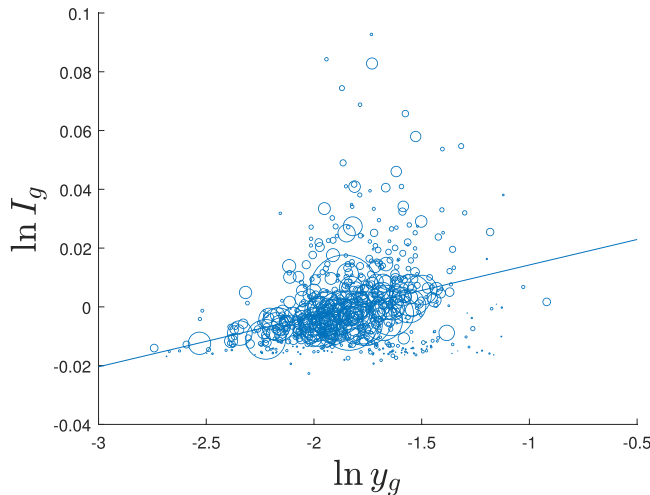


FIGURE A.11

Group-level import competition and income

Notes: The figure plots the relationship between $\ln I_g \equiv \ln \sum_s \pi_{igs} \frac{\beta_{is}}{\gamma_{is}}$, our measure for regional import-competition, and the logarithm of group-level average income per worker. The solid line displays the linear fit of this relationship, with each commuting zone weighted by its population size. The size of a circle indicates the population size of that commuting zone.

TABLE A.1
List of sectors

Sector Nr.	Sector description	β_s	r_s	β_s/r_s	$\lambda_{US,US,s}$
15–16	Food, Beverages, and Tobacco	0.03	0.03	0.98	0.95
17–19	Textiles and Textile or Leather Products	0.01	0.01	1.52	0.57
20	Wood and Products of Wood and Cork	0.01	0.01	1.09	0.86
21–22	Pulp, Paper, Printing, and Publishing	0.02	0.02	0.98	0.94
23	Coke, Refined Petroleum, and Nuclear Fuel	0.01	0.01	1.03	0.91
24	Chemicals and Chemical Products	0.02	0.02	1.01	0.82
25	Rubber and Plastics	0.01	0.01	1.01	0.89
26	Other Non-Metallic Mineral	0.01	0.01	1.06	0.85
27–28	Basic Metals and Fabricated Metal	0.03	0.02	1.06	0.86
29	Machinery, Nec	0.02	0.02	0.95	0.76
30–33	Electrical and Optical Equipment	0.04	0.04	1.07	0.63
34–35	Transport Equipment	0.04	0.03	1.06	0.73
36–37	Manufacturing, Nec; Recycling	0.01	0.01	1.26	0.67
	Non-manufacturing	0.75	0.76	0.98	0.98

Notes: This table lists the 14 sectors used in our analysis. The first column has the ISIC Rev.3 sectors for each of the manufacturing subsectors, and the second column has the sector description. The next three columns show the Cobb–Douglas expenditure share, the earnings share r_s and the sectoral import-competition index β_s/r_s for the US. The final column has the domestic expenditure share for the US, $\lambda_{US,US,s}$.

TABLE A.2
Fit of China shock counterfactuals to the data

	(1) Actual $\ln \hat{e}_g$	(2) Actual $\ln \hat{\pi}_{gHP}$	(3) Actual $\ln \hat{e}_g$	(4) Actual $\ln \hat{\pi}_{gHP}$
Model-predicted $\ln \hat{e}_g$ ($\alpha = 1/3$)	1.324 (1.734)			
Model-predicted $\ln \hat{\pi}_{gHP}$ ($\alpha = 1/3$)		2.047*** (0.562)		
Model-predicted $\ln \hat{e}_g$ ($\alpha = 1/2$)			1.925** (0.863)	
Model-predicted $\ln \hat{\pi}_{gHP}$ ($\alpha = 1/2$)				1.134*** (0.316)
Constant	−0.0808*** (0.00303)	−0.0465*** (0.00613)	−0.0829*** (0.00325)	−0.0515*** (0.00664)
Observations	722	722	722	722
R^2	0.001	0.032	0.007	0.014

Notes: ** $p < 0.05$, *** $p < 0.01$. Standard errors, in parentheses, are clustered at the state level. The table regresses values in the actual data for the period 2000–7 on simulated values for the counterfactual China shock. Columns 1 and 2 set $\alpha = 1/3$ in the simulations, while columns 3 and 4 set $\alpha = 1/2$.

TABLE A.3
Regression of predicted changes in employment and income per worker on ADH's China shock

	(1) $\ln \hat{\pi}_{gHP}$	(2) $\ln \hat{e}_g$	(3) $\ln \hat{y}_g$
$\sum_{s \in M} \pi_{gs} \Delta IP_{st}^{\text{China} \rightarrow \text{USA}}$	0.284*** (0.0834)	-0.0379*** (0.0111)	-0.0758*** (0.0222)
Mean and SD of dependent variable	-0.557 [0.0566]	0.0742 [0.0754]	0.148 [0.151]
First stage coeff.	0.528	0.528	0.528
First stage F	29.30	29.30	29.30
Observations	722	722	722
Instrument	China \rightarrow Other	China \rightarrow Other	China \rightarrow Other

Notes: *** $p < 0.01$. Standard errors, in parentheses, are clustered at the state level. The table regresses model predicted changes in employment and income per worker on the actual ADH's China shock for the period 2000–7. All dependent variables are based on Section 8's model and assume $\alpha = 1/3$ (for ease of exposition, all dependent variables are multiplied by 100. We also provide the means and standard deviations of the dependent variables in rows 3 and 4). The regressor of interest is the original ADH import exposure measure ($\sum_{s \in M} \pi_{gs} \Delta IP_{st}^{\text{China} \rightarrow \text{USA}}$), which is instrumented using the ADH instrumental variable: $\sum_{s \in M} \pi_{gst} \Delta IP_{st}^{\text{China} \rightarrow \text{Other}}$. All regressions include the same controls employed in ADH's preferred specification: lagged manufacturing shares, period fixed effects, Census division fixed effects, and beginning-of-period conditions (% college educated, % foreign-born, % employment among women, % employment in routine occupations, and the average offshorability index).

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Supplementary Data

Supplementary data are available at *Review of Economic Studies* online. And the replication packages are available at <https://dx.doi.org/10.5281/zenodo.6256415>.

Data Availability Statement

The data underlying this article are available in Zenodo at: <https://dx.doi.org/10.5281/zenodo.6256415>.

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