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Road Capacity, Domestic Trade and Regional Outcomes

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Road Capacity, Domestic Trade and Regional Outcomes

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Abstract

What is the impact on intra-national trade and regional economic outcomes when the quality and lane-capacity of an existing paved road network is expanded significantly? We investigate this question for the case of Turkey, which undertook a large-scale public investment in roads during the 2000s. Using spatially disaggregated data on road upgrades and domestic transactions, we estimate a large positive impact of reduced travel times on trade as well as local manufacturing employment and wages. A quantitative exercise using a workhorse model of spatial equilibrium implies heterogeneous effects across locations, with aggregate real income gains reaching 2-3 percent in the long-run. Reductions in travel times increased local employment-to-population ratio but had no effect on local population. We extend the model by endogenizing the labor supply decision to capture this finding. The model-implied elasticity of employment rates to travel time reductions captures about one-third of the empirical elasticity.

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1 Introduction

Uneven distribution of economic activity creates large and longstanding income inequality across regions within countries. High transport costs can cause spatial disparities in economic activity by impeding market access in isolated regions, both in terms of firms' ability to sell goods and in terms of their ability to buy the required inputs. To address this issue, governments around the world allocate large sums of money to transport infrastructure projects: developing countries invest around 1.5 percent, and the OECD countries around 0.8 percent of their gross domestic product in transport infrastructure annually (Kornejew, Maruyama Rentschler, and Hallegatte, 2019).

Investment in transport infrastructure can impact regional inequality and improve growth prospects by facilitating trade. But how large are these gains, especially when there are various types or stages of investments that are possible? Arguably, constructing a new road from scratch or paving a dirt road would have a different effect than constructing a highway or expanding the lane capacity of existing roads. Previous empirical work has focused on cross-country analysis (Limao and Venables, 2001; Yeaple and Golub, 2007), on the impact of the US interstate highway system (Duranton, Morrow, and Turner, 2014; Allen and Arkolakis, 2014; Jaworski and Kitchens, 2019), new road construction in the UK (Gibbons, Lyytikäinen, Overman, and Sanchis-Guarner, 2019), and the construction or paving of new roads in low- or lower-middle income countries, such as Faber (2014) on the highway network in China, Asturias, Ramos, and Santana (2018) on the Golden Quadrilateral highway in India, and Kebede (2019) on improved village roads in Ethiopia.

We complement the existing literature by offering new empirical and quantitative evidence on economic gains from investments in transport infrastructure. To the best of our knowledge, this is the first paper that quantifies the benefits of a major *capacity upgrade to existing transport infrastructure* in middle-income economies using rich spatially disaggregated data. In particular, we study the case of Turkey – a large upper middle-income country which undertook major public investments towards expanding the lane capacity of its existing highway network during the 2000s. Overall, the availability of high-quality domestic trade data within a geographically complex country, together with the possibility of exploiting a large-scale capacity upgrading of roads, provides a unique window into the study of economic gains from investment in transport infrastructure.

Our work differs from and contributes to the previous literature in two major ways: First, we use high-quality domestic *inter-district* trade data, which is generated by the universe of domestic firm-to-firm transactions. Domestic trade data between more than 900 Turkish districts is complemented with district-level employment and wage data over the

period 2006-2016. Using spatially disaggregated data helps our identification as we rely on variations in improvements in connectivity across district pairs within a province pair. We are also able to compute welfare gains at different regional levels, showing that welfare gains decrease with the level of aggregation, highlighting the quantitative role of variety gains from trade.

Second, canonical spatial models studying the impact of infrastructure improvements assume that population is fixed in the short-run while in the long-run, it can change through population re-allocation across space (Allen and Arkolakis, 2014; Donaldson and Hornbeck, 2016). We find that reductions in travel times increased employment-to-population ratio but had no visible effect on local population in the short-run, suggesting that shocks to market access can affect employment rates through other margins. To capture this finding, we extend a workhorse spatial model à la Allen and Arkolakis (2014) by endogenizing the labor supply decision through a standard consumption-leisure trade-off. We prove the existence of equilibrium in this extended model and derive the sufficient condition for it to be unique. Different from its analog in Allen and Arkolakis (2014), this sufficient condition highlights the interaction of agglomeration economies with labor force participation. The elasticity of employment with respect to travel time reductions implied by the calibrated model captures about one-third of the empirical elasticity.

In general, the main challenge faced by researchers when estimating the effects of transport improvements on key economic outcomes is to design a convincing empirical strategy: allocation of transport investment may not be random and might be correlated with other observed or unobserved location-specific factors. We address the potential endogeneity of the placement of transportation investment in a number of ways. First, as already discussed, we use highly spatially disaggregated data, which allows us to rely for identification on variations in connectivity improvements across district pairs within a province pair. We argue that, given the size of the districts—the median urban population is 10,400—and the national scale of the project, targeting district pairs would be unrealistic government micro-management, which lessens endogeneity concerns in this setting. Second, following Jaworski and Kitchens (2019) and Hornbeck and Rotemberg (2021), we exclude trade flows with nearby districts, i.e. districts located within the same province. This helps us avoid cases where the government targets certain provinces in its investment plan, allowing us to exploit improvements of a district's connectivity only with districts outside its own province. Third, as it already accounts for a significant share of the country's economic activity, we exclude the districts of Istanbul as source and destination. Finally, we control for initial bilateral distance between districts in our baseline specification to (i) ensure that our results do not simply reflect a mean reversion in connectivity, and (ii) account for route-dependent

trends. This is feasible thanks to the unique feature of Turkey's transportation investment program in 2000s, which resulted in substantial reductions in travel times between district pairs, leaving road bilateral distances between them (almost) unchanged.

The empirical exercise first measures the impact of road construction on reduced travel times, then links travel time reductions to changes in inter-district trade as well as local employment and wages. We leverage a new dataset on domestic trade across more than 900 districts in Turkey. The data span a time period during which intensive road construction took place (2006-2015) and can be broken down by industry to analyze heterogeneous effects as well as to control for compositional changes. The nature and the quality of data improves upon Coşar and Demir (2016) who have examined the effect of the same investment program on the external trade of Turkish provinces between 2003-2012 using provincial shares of upgraded roads in the road stock. In contrast, this paper uses district-to-district trade, which captures a larger fraction of total economic activity, and GIS-based district-to-district travel times, a more precise measure of transportation costs. To calculate bilateral travel time between Turkish districts, we have digitized the official maps of road network data before and after the investment program.

Our results suggest that travel time savings due to the investment program boosted domestic trade in Turkey. In particular, our preferred estimate implies that a 10 percent decrease in bilateral travel times from their initial average generates about 1 million USD increase in trade flows for a typical district pair over 10 years. The results are robust to a number of robustness checks, including a falsification test that investigates whether changes in domestic inter-district trade flows during the earlier years of the sample period (i.e. 2006-2011) can be explained by travel time reductions in the latter years (i.e. 2010-2015).

Next, we construct a variable that captures the average connectivity improvement for each district and examine how it affected local employment and wages. In the spirit of Borusyak and Hull (2020), we control for the initial connectivity of districts: while the authorities did not select investment locations according to their economic outcomes, more remote districts benefited more from road upgrades as the investment program was comprehensive. The estimated elasticity for local employment implies, for a district with the average connectivity in 2005, a 1-hour improvement in connectivity increases employment in goods-trading industries by 0.36 log points, which corresponds to close to half of the average district-level employment growth in those industries between 2006 and 2016. We also investigate the source of this increase by examining the responses of population and employment-to-population ratio to improvement in district-level connectivity. While we find no discernible effect on local population, the ratio of employment to population responds strongly to improvements in domestic transport connectivity. This result is similar to the

findings of Autor, Dorn, and Hanson (2013) and Dix-Carneiro and Kovak (2017) who document regional employment effects of aggregate trade shocks with no significant effects on population.

To gauge the welfare effects, we adapt a workhorse model of economic geography (Allen and Arkolakis, 2014) to the case at hand. The framework allows labor mobility within a standard Armington trade model, capturing the spatial equilibrium within a country in the long-run. We calibrate districts' productivities and amenities from their 2005 population shares and nominal wages. The quantified model helps us to calculate heterogeneous welfare changes across districts through market access shifts in the short run when labor is immobile: at conventional parameter values, the largest and median welfare gains across districts are 12.4 percent and 2.9 percent, respectively. In the long run, when labor is perfectly mobile across regions, the implied aggregate welfare increase is close to 3 percent. When we re-calibrate the model at higher levels of aggregation—81 provinces and 26 regions corresponding to Nomenclature of Territorial Units for Statistics (NUTS) levels 3 and 2, respectively— welfare gains fall slightly below 2 percent. This result highlights the importance of the variety channel in the model: given the Armington assumption, higher levels of spatial disaggregation implies more varieties, and thus a bigger welfare-enhancing role of improved market access.

To rationalize the empirical finding of increased local employment relative to population as a response to improvement in transport connectivity, we extend the model to incorporate an endogenous labor supply decision and then re-calibrate it. Using the model-implied short-run employment levels at fixed population shares, we replicate the reduced-form regression of the change in location-specific employment ratio on average travel time reductions. The results imply that the extended model is quantitatively relevant in explaining the reduced-form relationship between reduced travel times and employment rates: the model-implied elasticity captures about one-third of the reduced-form empirical elasticity.

The rest of the paper is organized as follows. Section 2 presents the country context and provides details of the transport investment program. Section 3 describes the data. In Section 4, we present the empirical strategy and results for the effects of road capacity improvements on domestic bilateral trade between Turkish districts. It is followed by the empirical results for local outcomes such as population, employment, and wages. Section 6 presents the model and calibration results for welfare, and Section 7 concludes the paper.

4

2 Background

Being an upper-middle-income country according to the World Bank classification, with a GDP per capita of USD 14,117 (in constant 2010 dollars) and a population of 79.8 million as of 2016, Turkey is a relevant setting to study the topic of interest: almost 35 percent of the world population live in the 60 upper-middle income countries which invest heavily in transport infrastructure, aiming to boost their future growth performance. Therefore, lessons learnt from Turkey's road upgrading program will be informative about other countries, particularly those at a similar level of development.

Turkey is administratively divided into 81 provinces which correspond to the NUTS 3 level in the Eurostat classification of regions. Each province is further divided into districts (*ilçe* in Turkish), which are lower level administrative divisions at the LAU-1 level in the EU classification. We use a consistent sample of 913 districts in our analysis.[1] Size and scale of districts correspond to those of communes, parishes or wards in other countries. District boundaries are shown with black lines in Figure 1, with urban town centers plotted with black dots within the boundaries.

Turkey is a large country with substantial internal trade costs: average bilateral distance between Turkish districts is almost 700 km, with a range from 55 km to 2200 km. This implies that transporting goods across different districts involves non-negligible costs. Roads constitute the dominant mode of domestic transportation in the country, accounting for about 90 percent of domestic freight (by tonne-km) and passenger traffic. This motivated the authorities to undertake a major public investment in its transportation infrastructure during the 2000s. The road network was already extensive prior to this investment: in 2005, a paved road network already connected Turkey's district centers (see thin grey lines in Panel A of Figure 1). However, the lack of dual carriageways for most network segments resulted in limited capacity, long considered inadequate (see the red lines indicating divided multi-lane highways or expressways).

Consequently, the Turkish government launched a large-scale transportation investment program in 2002 "to ensure the integrity of the national network and address capacity constraints that lead to road traffic accidents." The investment resulted in a significant percentage of existing single carriageways (undivided two-lane roads) being turned into dual carriageways.[2] By 2015, numerous arterial routes had been upgraded (see Panel B of Figure

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[1]During the sample period, the number of districts changed slightly from 913 to 922 due to boundary adjustments.

[2]According to the World Bank, Turkish public expenditures on transport have almost doubled from 1.06 percent of Gross Domestic Product (GDP) in 2004 to 1.92 percent in 2010, and the transport sector accounted for the bulk of the increase in total public investments over this period (http://bit.ly/2Aw0XX4).

1), with dual carriageways accounting for 35 percent of inter-provincial roads, up from 10 percent in 2002 (see Figure 2). The increase in capacity allowed vehicles to travel more reliably at higher speeds, making arrival times more predictable and reducing accident rates, with the number of fatalities per kilometre travelled declining by 57 percent between 2002-2014.[3]

Documentary evidence of the investment objectives and design do not suggest the selection of road segments for domestic trade-related outcomes. First, policy documents explicitly emphasize the long-term goal as the improvement of connections between all provincial centers to form a comprehensive grid network spanning the country, rather than boosting trade between particular regions. The General Directorate of Highways policy describes the criteria as "ensuring the integrity of the international and national networks, and addressing capacity constraints that lead to road traffic accidents."(GDH, 2014). Second, the extent of road upgrading shows considerable variation across provinces, without any visible sign of concentration in particular regions. Finally, the investment was centrally planned and financed from the central government's budget with no direct involvement of local administrations.[4]

In our analysis, we pursue a demanding identification strategy and exploit variations in improvements in connectivity across districts within a province pair. As this strategy already conditions on province pairs, it provides a plausibly clean identification of the trade effects of improvements in connectivity. While the authorities may target some province pairs in their investment plan, it is unlikely that they would target a particular district pair. As we will explain in detail below, we also take additional measures to alleviate concerns about other possible sources of omitted variable and selection biases.

3   Data

A distinguishing feature of our study is the availability of high-quality data on domestic trade flows within Turkey during a time period when the country undertook a significant upgrading of its road network. The source of the domestic trade data is the administrative firm-to-firm transaction data provided by the Turkish Ministry of Industry and Technology (MoIT, henceforth). Since 2006, Turkish firms have been legally required to report all purchases and sales transactions with a given business partner if the annualized amount of purchases/sales

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[3]See the second column from right in table 1 in Murat and Zorlu (2018). Since the reporting criteria was changed in 2015 from "fatality on impact" to "fatality within 30 days of the accident," we report the change until 2014.

[4]Additional details about the investment program and discussion of external evidence on its contribution to the improvement of road transport quality in Turkey are available in Coşar and Demir (2016).

exceeds a certain threshold ($\approx$USD 3,300 in 2010) to the Ministry of Finance.[5] The objective of this requirement is to reduce tax evasion and increase value-added tax (VAT) collection. Each transaction report is cross-checked and in case of inconsistencies, both firms are audited to retrieve the correct information.[6]

In this paper, we use annual bilateral trade flows in goods between Turkish districts constructed by aggregating the domestic firm-to-firm trade data described above. To study longer term effects of improvements in road capacity, we focus on changes in the outcome variables over a 10-year period between 2006-2016. We focus on domestic sales of firms that operate in manufacturing and wholesale industries.[7] The underlying micro-level data are generated by about 500 thousand buyers and sellers located in 913 Turkish districts.[8]

To measure the impact of the road upgrades, we calculate the decadal change in inter-district travel times. To do so, we digitized the official maps of the road network published by the General Directorate of Highways for 2005, 2010 and 2015. Figure 1 shows the first and last year's rendered maps. Using geographic information system (GIS) software, we then calculated the fastest possible travel times between the 913 provincial centers in each year. The maps in Figure 1 show both divided expressways and highways as dual carriageways. Since congestion is not an issue for these high-capacity roads, we base navigation speeds on tolerated margins over speed limits and assume a speed of 90 km/h on expressways and 110 km/h on highways. On single carriageways, where congestion is common, the average truck speed—as measured in a representative sample of non-highway road segments by the General Directorate of Highways[9]—was 72 km/h in 2010 when the share of expressways in the total road stock was 29 percent (Figure 2). To match this average, we solve $90 \cdot 0.29 + X \cdot (1 - 0.29) = 72$ for X and impute a 65 km/h average speed on single carriageways. For each pair of provincial centers in Figure 1, ArcMap software is used to calculate the shortest possible

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[5]We only include in the sample trade flows between domestic firms, i.e. export and import transactions are excluded.

[6]The micro data has been used in a recent study by Demir, Fieler, Xu, and Yang (2021). There is a growing empirical literature that uses domestic VAT data: e.g., Bernard, Dhyne, Magerman, Manova, and Moxnes (2019) and Tintelnot, Kikkawa, Mogstad, and Dhyne (2018) use Belgian VAT data; Huneeus (2018) uses Chilean data; Alfaro-Urena, Manelici, and Vasquez (2020) use Costa Rican data; Adao, Carrillo, Costinot, Donaldson, and Pomeranz (2020) use data from Ecuador.

[7]The agricultural sector is excluded since it is dominated by unincorporated small farmers whose transactions tend to fall under the reporting threshold. On the buyer side, we include all industries reported in the data. This excludes most utilities, all public services, and financial services which are not subject to VAT.

[8]Since firm-to-firm trade flows are recorded at the firm rather than at the establishment level, transactions of multi-establishment firms are accounted for at the headquarter location. The ensuing mismeasurement is most severe in utilities and financial services with numerous bank branches.

[9]This data is available at https://www.kgm.gov.tr/Sayfalar/KGM/SiteTr/Istatistikler/TrafikveUlasim.aspx

travel time for both years on the basis of the above assumptions regarding speeds.[10]

The average travel time between any two districts has been reduced by 1.4 hours, relative to the average of 9.8 hours in 2005. Time savings increase the further apart districts are, reaching three hours, on average, in the case of districts that are 1,500 km apart. While there is a significant negative correlation between reduction in travel times between districts and their initial distance, bilateral distance is not a strong predictor of changes in travel times. A regression of the reduction in travel times on distance yields an R^2 of 0.09. Figure 3 shows a binned scatter plot of travel time reductions from 2005 to 2015 against time-invariant bilateral distances.

Table 1 shows that an average Turkish district was buying from 40 districts and selling to about 50 districts in 2006. Both increased by about three times in 2016. Interestingly, the average value traded did not change between the beginning and end of the sample period. This could be a result of relatively smaller values traded for newly established trade links in 2016. For already existing trade relationships, the value of trade increased, on average, by 1.3 percent.

In the second part of the empirical analysis, we investigate the effect of road upgrades on district-level outcomes such as population, employment, and wage per worker. District-level population data come from the Social Security Institute (SSI). Data on district-level formal employment and wages are based on SSI administrative records and made available by MoIT. In the analysis below, we aggregate employment records and wage payments of establishments at the level of districts. We deflate wage payments using the consumer price index (2003=100).

As presented in Figure B1, our data closely replicate the country-wide aggregate formal employment numbers in private sector as well as their provincial distribution in 2011 and 2016. When compared to the publicly reported employment statistics of SSI, our dataset slightly undershoots the provincial employment figures for two reasons. First, we use the average of quarterly employment records of private establishments while the publicly available SSI data report the end of the year records. Second, our data exclude some industries,

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[10]We recalculated the optimal routes and resulting bilateral travel times between Turkish *provinces* under two alternative scenarios. Both scenarios keep the average speed on highways the same as in the baseline. Under the first scenario, we assume that the average speed on four-lane expressways is lower at 80 km per hour compared to 90 km per hour under the baseline. The second scenario, on the other hand, assumes a higher speed for single carriageways (75 km/hr) than the baseline (65 km/hr). The upper panel of Figure B4 plots changes in bilateral travel times under the two alternative scenarios against the baseline. For both scenarios, travel time changes are highly correlated with those under the baseline.

namely mining, utilities, public services, and financial services.[11]

4 Road Capacity and Domestic Trade

4.1 Baseline Results

We start our analysis by checking whether the reduced travel times resulting from the road improvements between 2005 and 2015 increased bilateral domestic trade flows between Turkish districts. Aggregating the data up to the level of district pairs, there are 833,569 pairs (913×913) that can potentially trade with each other as buyers or sellers. However the domestic trade matrix is highly sparse: as presented in Table 1, an average Turkish district was trading with less than 5% of potential partners in 2006. While there was a substantial increase in the average number of trading partners by the end of the sample period, the fraction of positive trade links was still less than 15% of the number of potential links. Below, we will present additional results that account for this sizable extensive margin increase.

Our main outcome variable is the 10-year change (between the initial and terminal years of the 2006-2016 period) in the logarithm of the value of bilateral trade between source district i and destination district j. Letting

$$\Delta \ln TravelTime_{ij} \; = \; \ln(TravelTime_{ij}^{2015}/TravelTime_{ij}^{2005}),$$

we estimate

$$\Delta \ln Trade_{ij} \; = \; \theta_1 \cdot \Delta \ln TravelTime_{ij} + \theta_2 \cdot \ln Distance_{ij} + \phi_i + \phi_j + \phi_{p_i p_j} + \epsilon_{ij}, \qquad (1)$$

where source and destination district fixed effects (ϕ_i and ϕ_j) control for district-level characteristics that affect domestic sales and purchases of each district, e.g., productivity improvements, changes in population, etc. They also account for initial location advantages of buyer and supplier districts, such as centrality in the transport network. Since road upgrades led to significant travel time savings, we expect $\theta_1 < 0$. We use two-way clustered standard errors by source and destination districts.

To alleviate possible omitted variable and selection biases, we take the following measures. First, to account for the possibility that the government targeted particular province

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[11]SSI does not publish data for the year 2006, thus we cannot include a comparison for the initial year of the sample period. However, bearing in mind the exclusions noted in the text, the nationwide formal employment in private sector in 2007 as reported by SSI (8.5 million workers) is not far from the aggregate number calculated from our dataset for the year 2006 (7 million).

pairs in the investment plan, we add province-pair fixed effects, $\phi_{p_i p_j}$. Given the size of the districts and the national scale of the project, to target district pairs would be micro-management by the government to an unrealistic degree, and therefore, is not a concern in this setting. Therefore, we identify the effect of travel time reductions on trade from variations across district pairs within a given province pair. Even for district pairs that are located (almost) equidistant from each other for a given province pair, the shortest routes might pass through different connections. Figure 4 shows an example. Consider a province pair, Adana and Konya, each of which is divided into multiple districts. Panel A shows the shortest route between Bozkir (Konya) and Ceyhan (Adana). These two districts are located 429 km apart, and the travel time between them is 284 minutes (4 hours and 44 minutes), implying an average speed of 90.6 km per hour. Panel B shows the shortest route between two other district pairs for Konya and Adana, namely Seydisehir and Aladag. The shortest route between them is almost the same as the one between Bozkir and Ceyhan. However the average speed on the route is only 71.3 km per hour as the overlap of road segments between the two routes is very small. These two district pairs for Konya-Adana benefited quite differently from road improvements in the second half of 2000s. While the shortest travel time between Bozkir and Ceyhan decreased by 39 minutes (12 percent of initial travel time), the reduction of travel time between Seydisehir and Aladag was 98 minutes (22 percent of initial travel time). This is the variation we will exploit in our empirical analysis.

Second, we exclude trade flows between district pairs within the same province. This restriction rules out the possibility that our results are driven by significant road improvements in particular provinces, which led to an increase in within-province trade. Third, we exclude the districts of Istanbul as source and destination. Istanbul was already well connected to the major industrial clusters of the country in the beginning of the period. Since the region accounts for a significant share of economic activity, it is large enough to create a bias in our estimates. Fourth, on the supply side, we exclude those provinces which already had highway connections in the beginning of the sample period. Those highways are represented with red lines in Panel A of Figure 1. The excluded provinces are Ankara, Izmir, and Mersin. Except for Ankara (the capital city), these provinces are the major international ports of the country. Finally, we include initial bilateral distance between districts as a control variable to (i) ensure that our results do not simply reflect a mean reversion in connectivity, and (ii) account for geography-dependent trends.

Table 2 reports the baseline results. In the first column, we keep Ankara, Izmir, and Mersin in the sample of source districts. The coefficient on $\Delta \ln TravelTime_{ij}$ has the expected sign and is estimated to be economically and statistically significant. Results obtained for our preferred specification are presented in column 2, where districts that belong to the

provinces with highway connections in 2005 are excluded from the sample of suppliers. The estimate for the variable of interest slightly increases in magnitude compared to the first column.[12] The estimate implies that a one-hour reduction in travel time between two districts, which corresponds to a 10 percent decrease at the mean travel time in 2005, increases bilateral trade between those districts by around 8.2 percent. This effect is statistically significant and translates into an almost 1 million USD increase in trade flows for a typical supplier district over 10 years.[13]

There exist two channels through which improvements in domestic transport infrastructure affect inter-provincial trade: first, by reducing the cost of transporting goods between the source and destination provinces, and second, by reducing the cost of finding new suppliers/buyers (i.e. establishing new trade relationships, or the extensive margin of trade) as in Bernard, Moxnes, and Saito (2019). To further examine the extensive margin effect of reduced travel times on the establishment of new trade links, we estimate a linear probability model in which the dependent variable equals 1 for province pairs with positive trade in 2016 conditional on zero trade in 2006, and 0 otherwise. The result in column (3) of Table 2 suggests a district pair that experienced a 10 percent decline in travel time had a probability of 11 percent to start trading in 2016. The last column combines the intensive and extensive margin effects by replacing the dependent variable from equation (1) with the hyperbolic sine transformation of trade flows, which keeps zero flows in the estimation sample. The estimated coefficient on $\Delta \ln TravelTime_{ij}$ almost doubles compared to the estimate in column (2).

4.2 Robustness Checks

We subject the baseline results to three robustness checks. The first involves splitting the sample into sub-periods and estimating a placebo test. Table B1 presents the results from this placebo test which regresses changes in trade flows in the 2006-2011 period (first half of the sample period) on travel time reductions in the succeeding 2010-2015 period (second half of the sample period). If our baseline results reflect some general trend in trade flows driven by omitted factors correlated with road construction, then changes in bilateral trade in 2006-2011 could also be explained by road improvements in the succeeding period. As expected, improvements in the succeeding period are statistically and economically insignificant, which strengthens the validity of our identification.

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[12]Figure B3 presents predicted and actual values of changes in bilateral trade flows between districts.

[13]Typical district is defined as follows: (i) its number of buyer districts is equal to the average number (52) across all districts in 2005, and (ii) its travel time from its buyer districts is equal to the average (602.6 minutes) across all district pairs in 2005.

The underlying micro-level VAT data are recorded at the level of firms rather than establishments. This implies that transactions by multi-establishment firms, which operate in multiple locations, are all accounted for at the headquarter location. This creates a measurement error that can potentially affect our estimates. Therefore, in another robustness check, we exclude multi-establishment firms as suppliers and buyers before aggregating trade flows at the level of district pairs. As presented in column (2) of Table B1, the estimate for the intensive margin of trade obtained from this restricted sample is significantly larger than our baseline estimate (column (2) of Table 2). One possible explanation for this difference is that multi-establishment firms are generally larger so they can already pay the fixed cost of reaching relatively distant markets. Therefore, it is not surprising to see larger effects when these firms are excluded from the sample. However, when both the intensive and extensive margin effects are taken into account in column (3), the estimate obtained for the variable of interest is very close to the baseline estimate reported in column (4) of Table 2.

Finally, Figure B2 presents the distribution of the estimate of θ_1 in equation (1) on 500 randomly drawn samples of district pairs, each of which has a size that is about one-third of the baseline sample. The aim of this exercise is to test whether the estimates are sensitive to sample selection. We conduct this test for the intensive margin of trade (column 2 of Table 2) as well as the combined intensive and extensive margins (column 4). Regardless of the specification of the dependent variable, 75 percent of the estimates obtained from the randomly drawn small samples fall within the confidence interval obtained for the respective full-sample estimates. Despite a considerable reduction in sample size, θ_1's are precisely estimated: 80% of the estimates obtained from randomly sampled district pairs are significant at conventional levels for the intensive margin of trade. The fraction increases to 96% for the combined intensive and extensive margins. This robustness check alleviates potential concerns about dominance of certain district pairs, as well as selection of the location of road upgrades by the authorities.

4.3   Industry Heterogeneity

Next, we use the industry dimension of the data to investigate whether trade effects of road upgrading vary across industries. More transportation-intensive industries, based on, for instance, time sensitivity as in Hummels and Schaur (2013) or heaviness as in Duranton, Morrow, and Turner (2014), could benefit more from road upgrades when selling their goods longer distances.

To estimate industry-level heterogeneity in terms of transportation intensity, we restrict the sample to manufacturing industries (i.e. 2-digit NACE codes) and augment the

specification in (1) by interactions with two industry characteristics: $TimeSensitivity$ and $Heavy$. The former refers to the share of trade value transported by air in an industry, and the latter captures the (logarithm) of the weight to value ratio. Both measures are constructed by Coşar and Demir (2016) using 2005 UK import data. In addition to source and destination district fixed effects as in equation (1), we include province pair fixed effects specific to each 2-digit NACE industry code. This rich set of fixed effects absorb, among others, the intensity of pre-existing bilateral trade flows at industry level.

The results are presented in Table 3. In the first column, we exclude interactions between bilateral distance and industry characteristics, and add them in the second column. In the last column, we run a demanding specification with origin-destination pair fixed effects. In this specification, bilateral distance and change in bilateral travel time are absorbed by the pair fixed effects and cannot be identified. However, we can still identify the interaction terms from variations across industries within district pairs. While the estimated coefficient on the interaction between travel time changes and time sensitivity $(TimeSensitivity_s * \Delta \ln TravelTime_{ij})$ is statistically insignificant at conventional levels in all columns, the estimate for the interaction with heaviness remains significant and robust to various specifications. Importantly, the estimated coefficient on $\Delta \ln TravelTime_{ij}$ is insignificant throughout. Instead, the effect of reduced travel times works through increased trade between districts in industries that are characterized by high weight-to-value ratio. This suggests that transportation intensity of goods in terms of heaviness is an important channel through which travel time savings due to road upgrades increase domestic trade flows.

5   Road Capacity and Regional Outcomes

Beyond its impact on trade, did the reduction in domestic travel times affect other key regional economic outcomes such as employment and total sales? To address this question, we construct a variable capturing improved domestic market access at the district level. In particular, weighting each district's time savings on the basis of destination districts' population for 2005,

$$\Delta \ln TravelTime_i^{WgtAvg} \ = \ \sum_{j=1}^{N_i} \left(\frac{population_j}{\Sigma_{k=1}^{N} population_k} \right) \cdot \Delta \ln TravelTime_{ij}, \tag{2}$$

calculates the average connectivity improvement experienced by a district when selling goods to other districts. In the formula, the total number of districts in the denominator (N) is

913. To alleviate endogeneity concerns and de-emphasize the role of initial concentration of economic activity, we exclude destination districts within the same province as the origin, thus the upper bound of the summation index (N_i) is district specific. We also construct an unweighted average of changes in bilateral travel times for each district:

$$\Delta \ln TravelTime_i^{Avg} = \sum_{j=1}^{N_i} (1/N_i) \Delta \ln TravelTime_{ij}. \tag{3}$$

While the variable defined in equation (2) takes the initial location advantages (or disadvantages) of a district seriously by assigning a larger weight to travel time changes with respect to larger destinations in terms of population, the one defined in (3) treats each destination district symmetrically.

We will report the results from estimating equations of the form

$$\Delta Outcome_i = \phi_{p_i} + \mu_1 \cdot \Delta \ln TravelTime_i^{WgtAvg} + \mu_2 \cdot \ln Dist_i^{WgtAvg} + \epsilon_i, \tag{4}$$

where $Outcome_i$ is ratio of employment to population, logarithms of district-level employment, population or wage payments per worker depending on the specification. Aggregate province-level changes are controlled by including ϕ_{p_i}. $\ln Dist_i^{WgtAvg}$, i.e. population-weighted average of bilateral distance at the district level, further controls for the initial connectivity of districts. As in the previous section, we exclude from the estimation sample the four provinces with highways in the beginning of the period, namely Ankara, Istanbul, Izmir, and Mersin. Standard errors are clustered at the province level.

Table B2 presents how the variable of interest, $\Delta \ln TravelTime^{WgtAvg}$, correlates with initial district-level characteristics such as population density and employment-to-population ratio. The first two columns do not control for variation across provinces. According to column (1), districts with lower population density and employment experienced larger improvements in average connectivity during the 2005-2015 period. However, the size of both estimates drops substantially when average initial bilateral distance, $\ln Dist^{WgtAvg}$, is added as a control variable in column (2). This result echoes what we argued before: while the authorities did not select investment locations based on economic outcomes, more remote districts benefited more from road upgrades. This highlights the importance of controlling initial remoteness of districts in equation (4). In column (6), we add province-level fixed effects. Conditional on province-level fixed effects, (weighted or unweighted) average change in bilateral travel times does not correlate with initial district-level characteristics, including remoteness of districts. These results increase our confidence in the validity of our identification strategy.

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In a recent paper, Borusyak and Hull (2020) show a new source of omitted variable bias in empirical settings similar to ours. They argue that geographical units that have a central place in the transportation network would benefit more from construction of new roads even under the assumption of randomly placed road improvements. To correct such bias, the authors suggest re-centering market access growth using a measure of "expected growth of market access". To illustrate the importance of re-centering, they focus on the construction of high-speed rail network in China between 2007-2016: while the distribution of actual market access changes is highly geographically clustered, re-centering successfully removes the clusters.

Upper panel of Figure 5 shows the spatial distribution of $\Delta \ln TravelTime^{WgtAvg}$ across Turkish districts. As in the case of China's railroad construction, the improvements in district-level travel times are highly concentrated. In the Turkish setting, road construction was comprehensive. Therefore, the concentration reflects the initial remoteness of districts: districts located in the eastern part of the country, which were less connected to the more economically active parts of the country before the start of the investment program, benefited disproportionately more from road upgrades. Fixed effects and the control variable included in equation (4) serve to correct this bias. To show this, we regress our variable of interest on province-level fixed effects and district-level population-weighted average of initial distance ($\ln Dist_i^{WgtAvg}$) and obtain the residuals. Panel B of Figure 5 plots the distribution of those residuals. As opposed to the upper panel, adjusted average of travel time changes do not show any geographical concentration, and many districts change categories depending on their initial remoteness.

The results from estimating equation (4) are presented in Table 4. The outcome of interest is total employment in the first column. The coefficient on $\Delta \ln TravelTime_i^{WgtAvg}$ is estimated to be negative and statistically significant. When we focus on employment in manufacturing (and wholesale) in column (2), the parameter of interest becomes even larger, implying that improvements in domestic market access have a positive effect on aggregate manufacturing employment. The estimated elasticity of 3.4 implies, for a district with the average connectivity in 2005, a 1-hour improvement in connectivity increases employment in goods-trading industries by 0.36 log points, which corresponds to slightly more than half of the average district-level employment growth in those industries between 2006 and 2016.[14]

We subject this result to two robustness checks. In column (3) of Table 4, we replace

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[14]In our data, average value of the population-weighted bilateral travel time is 9.6 hours, and the average employment growth in manufacturing and wholesale industries across districts is 63 percent. Note that the average growth across more than 900 districts is much higher than the 30 percent aggregate formal employment growth in these industries over the same time period.

population-weighted changes in bilateral travel times with the simple average as defined in equation (3). This variable measures district-level improvements in connectivity by treating all districts equally. The estimated elasticity from this specification is even larger than the baseline estimate. In column (4), we run a placebo test where we regress employment changes between 2006-2011 on population-weighted changes in bilateral travel times between the second half of the period, i.e. 2010-2015. As expected, the estimate of the coefficient of interest shrinks, changes sign, and loses statistical significance. This test increases our confidence that our baseline results do not capture increased economic activity due to planned investment in road infrastructure or some form of pre-trends.

In the last two columns of Table 4, we investigate whether our finding that improvements in domestic transport connectivity lead to local employment growth is driven by rising population in those regions. In column (5), we use district-level population growth as an outcome variable in specification (4). While the coefficient of interest is estimated to be negative, it is not statistically significant at conventional levels. In the last column, we estimate the effect of improvements in domestic transport connectivity on the ratio of local employment to population. Consistent with the earlier results, ratio of total employment to population responds strongly to improvements in domestic transport connectivity.[15]

Finally, in Table 5, we estimate the wage effects of road upgrades at the district level. The dependent variable is total wage payments divided by total employment in the first column. The estimated coefficient on $\Delta \ln TravelTime^{WgtAvg}$ is negative and economically meaningful. Although the coefficient estimate becomes smaller, the result is robust to constructing district-level wage per worker using employment records of manufacturing and wholesale establishments (column 2). As presented in the last column, the result that improved domestic transport connectivity raises real wage per worker is not driven by expected future investments.

While our results so far suggest substantial regional effects, one cannot aggregate estimated local impacts due to treatment spillover effects between districts. Moreover, counterfactual statements about real income necessitate the construction of a theory-based price index and incorporation of labor reallocation in the long-run. To do so, the next section presents and calibrates a workhorse spatial equilibrium model with trade.

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[15]We highlight that the data on local employment responses do not come from a labor force survey because Turkish regional labor statistics based on household labor surveys are only available at an aggregate level of 26 NUTS-2 regions. Therefore, we calculate districts' employment to population ratios by simply adding local formal employment in industries included in our analysis and dividing it by official urban populations.

6 Quantitative Analysis

The baseline model of economic geography and trade closely follows Allen and Arkolakis (2014). We first describe the environment and long-run spatial equilibrium. Quantitative applications and welfare calculations based on the calibrated model distinguish short- and long-run outcomes when population is fixed and mobile, respectively. We then extend the model to feature endogenous local labor supply, motivated by our finding of increased local employment in the absence of spatial population reallocation. After establishing sufficient conditions for the existence and uniqueness of equilibrium in the extended model, and calibrating its additional parameters, we compare model-generated outcomes with relevant empirical findings as well as with the baseline model.

6.1 Model

An exogenous aggregate labor supply, normalized to $\overline{L} = 1$ is freely mobile between the N locations of the country. Each location i produces a differentiated Armington variety under perfect competition. The cost of trade between two locations j, i is of iceberg type: $T_{ij} = T_{ji} > 1$ if $i \neq j$, and $T_{ii} = 1$. That is, location-i variety with an origin price p_i costs $p_{ij} = T_{ij} p_i$ in location j. Production is competitive and linear in labor with location-specific productivity:

$$A_i = \overline{A}_i L_i^\alpha. \tag{5}$$

The exogenous component $\overline{A}_i$ is augmented by population so that production displays external increasing returns to scale due to agglomeration forces if $\alpha > 0$. Given a local nominal wage of w_i, competitive linear production implies an origin price of $p_i = w_i / A_i$.

Similarly, each location has an exogenous amenity level $\overline{u}_i$, augmented by its population:

$$u_i = \overline{u}_i L_i^\beta. \tag{6}$$

Amenities display decreasing returns to scale due to congestion forces if $\beta < 0$.

Household utility features two parts: consumption and amenities. Consumption preferences over differentiated varieties are CES with an elasticity of substitution $\sigma \in (1, \infty)$, which implies a price index of $P_i = (\sum_{j=1}^N p_{ji}^{1-\sigma})^{1/(1-\sigma)}$. The utility of a representative household living in location i is given by

$$W_i = \frac{w_i}{P_i} \cdot u_i. \tag{7}$$

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CES demand implies trade flows from j to i equal to

$$X_{ji} = p_{ji} q_{ji} L_i = \left(\frac{T_{ji} p_j}{P_i}\right)^{1-\sigma} w_i L_i, \tag{8}$$

where q_{ji} is variety-j consumption by a representative household residing in location i.

Long-run spatial equilibrium holds when wages and population allocation $\{w_i, L_i\}_{i=1}^N$ are such that product markets clear: $q_i = A_i L_i = \sum_{k=1}^N T_{ik} q_{ik}$; each location's expenditures equal its total income: $\sum_{j=1}^N X_{ij} = w_i L_i$; welfare is equalized across locations: $W_i^* = W$ and aggregate population constraint holds: $\sum_{i=1}^N L_i = 1$.

Allen and Arkolakis (2014) characterize the conditions on parameters that ensure the existence of an equilibrium. In particular, regardless of the magnitude of σ, a unique and stable equilibrium exists if $\alpha + \beta \leq 0$. Under this assumption, which we maintain, there is a one-to-one relationships between the set of exogenous productivities and amenities, $\{\overline{A}_i, \overline{u}_i\}$, and the set of endogenous wage and population levels $\{w_i, L_i\}$. Thus, given the empirical levels of $\{w_i, L_i\}$ and the function of trade costs between locations $T_{ji}^{1-\sigma}$, the following system of equations can be solved to back out composite amenities $u_i^{1-\sigma}$ and productivities $A_i^{1-\sigma}$ up to a scale W:

$$u_i^{1-\sigma} = W^{1-\sigma} \sum_{j=1}^N T_{ji}^{1-\sigma} w_i^{\sigma-1} w_j^{1-\sigma} \cdot A_j^{\sigma-1}, \tag{9}$$

and

$$A_i^{1-\sigma} = W^{1-\sigma} \sum_{j=1}^N T_{ji}^{1-\sigma} L_i^{-1} w_i^{-\sigma} L_j w_j^{\sigma} \cdot u_j^{\sigma-1}. \tag{10}$$

With values of $\{A_i^{1-\sigma}, u_i^{1-\sigma}\}$ at hand, exogenous components $\{\overline{A}_i, \overline{u}_i\}$ can be backed out for given values of (α, β, σ).

To calculate the level of welfare W and labor allocations $\{L_i\}$ implied by any vector of trade costs T_{ji}—for instance when travel times are reduced due to road upgrades—we keep exogenous productivities and amenities fixed at their calibrated values, and solve the following set of N equations together with the national labor market clearance condition:

$$L_i^{\tilde{\sigma}\gamma_1} = W^{(1-\sigma)} \overline{u}_i^{(1-\tilde{\sigma})(\sigma-1)} \overline{A}_i^{\tilde{\sigma}(\sigma-1)} \sum_j^N T_{ji}^{(1-\sigma)} \overline{A}_j^{(1-\tilde{\sigma})(\sigma-1)} \overline{u}_j^{\tilde{\sigma}(\sigma-1)} L_j^{\tilde{\sigma}\gamma_2}. \tag{11}$$

Here, $(\tilde{\sigma}, \gamma_1, \gamma_2)$ are functions of the parameters (σ, α, β).[16] Equation (11) follows from spatial utility equalization. We refer the reader to Allen and Arkolakis (2014) for the proofs and

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[16]In particular, $\gamma_1 = 1 - \alpha(\sigma-1) - \beta\sigma$, $\gamma_2 = 1 + \alpha\sigma + (\sigma-1)\beta$ and $\tilde{\sigma} = (\sigma-1)/(2\sigma-1)$.

the description of the solution algorithm.

6.2 Calibration and Quantitative Results

We calibrate the baseline model at the district level, starting with the estimation of the structural gravity equation (8) under the following specification:

$$\ln X_{ij} = \mu_i + \mu_j + (1-\sigma)\ln T_{ij} + \epsilon_{ij}, \tag{12}$$

where μ's are origin and destination fixed effects capturing location-specific terms. We specify trade costs as a function of travel times, $T_{ij} = TravelTime_{ij}^{\zeta}$. Substituting this into equation (12) :

$$\ln X_{ij} = \mu_i + \mu_j + \underbrace{(1-\sigma)\zeta}_{\delta} \cdot \ln(TravelTime_{ij}) + \epsilon_{ij}. \tag{13}$$

As standard in the literature, this estimation cannot separately identify the elasticity of trade to trade costs ($\sigma - 1$) from the elasticity of trade costs to travel times ζ. Trade costs, however, only appear as $T_{ij}^{1-\sigma}$ in the model. Hence, for the purpose of pinning them down, we do not need to take a stand on σ. We estimate equation (13) using both 2006 and 2016 trade flows between districts and report in Table 6. Columns (1)-(2) present OLS estimates while last two columns present weighted least squares estimates, with trade values used as weights. The latter is known to yield similar estimates to PPML (Mayer, Vicard, and Zignago, 2019), which proves to be infeasible in our setting due to the high number of fixed effects. We pick the mid-value of WLS coefficients from columns (3)-(4) equal to $\hat{\delta} = -0.81$. This provides us with the trade cost matrix used in quantitative exercises, $\hat{T}_{ij}^{1-\sigma} = TravelTime_{ij}^{\hat{\delta}}$, where $TravelTime_{ij}$ is either at its 2005 or 2015 value, normalized by imputed time-invariant within-district travel time.[17]

To calibrate β, the parameter capturing congestion forces, we use the isomorphism of the model to one that features residential land/housing in consumption (Allen and Arkolakis, 2014). In that version of the model, the price of the immobile fixed factor (land) is increasing in population, thereby decreasing the utility of residents. The isomorphism holds if land has a Cobb-Douglas expenditure share of $-\beta/(1-\beta)$. According to the Household Budget Survey of the Turkish Statistical Institute, housing has a stable expenditure share around

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[17]Our travel time data does not give us within district travel times. We impute it as follows: using the method of Leamer (1997), the internal distance for the average-sized district is $\sqrt{804/\pi} = 16$ km. At a speed of 65 km/h, this implies a travel time of 15 minutes within districts. We winsorize all travel times from below by 15 minutes. Finally we normalize them by that level so that $T_{ii} = 1$.

25 percent across the relevant data period.[18] We set $\beta = -1/3$ to match that value, which is very close to the value of $\beta = -0.3$ in Allen and Arkolakis (2014) who use the US housing expenditure share as the calibration target.

Estimates of the agglomeration parameter α in the literature range between 0.04 and 0.1 (Rosenthal and Strange, 2004). We use the mid-point of this range at 0.07 which satisfies the constraint $\alpha \in [0, -\beta]$ for the existence and uniqueness of equilibrium. For the elasticity of substitution, we take a baseline value of $\sigma = 5$ to attain a trade elasticity of $\sigma - 1 = 4$ (Simonovska and Waugh, 2014). Table 7 summarizes location-invariant parameter values and their calibration.

Finally, we calibrate location-specific parameters $\{\overline{A}_i, \overline{u}_i\}$ using district-level wages and populations $\{w_i, L_i\}$ as of 2005. Since Turkish Statistical Institute does not publish income per capita at this level of geographic disaggregation, we calculate districts' average wages from our data, dividing total wage bill by formal employment. Since the sectors used in the analysis (see footnote 7) are primarily urban, we use districts' urban population as reported by the Statistical Institute.

The calibration procedure is as follows: given initial trade costs $\hat{T}_{ij}^{1-\sigma} = TravelTime_{ij}^{\hat{\delta}}$ using 2005 travel times, and the 2005 empirical levels of $\{w_i, L_i\}$, we solve the system (9)-(10) for $\{A_i, u_i\}$ at the baseline level of welfare $W = 1$. With composite productivities and amenities at hand, we then back out $\overline{A}_i = A_i L_i^{-\alpha}$ and $\overline{u}_i = u_i L_i^{-\beta}$. In Figure B5, we plot calibrated values of $\{\overline{A}_i, \overline{u}_i\}$ against the data from which they were backed out. Evidently, productivities strongly correlate with nominal wages while amenities explain a larger share of the variation in population.

Short- and Long-run Outcomes   When labor is immobile, road upgrades generate spatial inequality between districts through changes in market access. To solve for the short-run equilibrium, we keep the population vector $\{L_i\}$ in its 2005 level, change trade costs T_{ij} to its lower level at 2015 travel times, and find market clearing wages w_i for each district. We then calculate district-level price indices P_i using the lower trade costs. Since labor is fixed, amenities enjoyed by residents do not change. The only variation in welfare comes from the real wage component of utility, that is, from the response of w_i/P_i to the change in trade

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[18]We use the 2005-2015 median values of housing and rent share in household consumption expenditures across regions as reported by the Turkish Statistical Institute: https://data.tuik.gov.tr/Kategori/GetKategori?p=gelir-yasam-tuketim-ve-yoksulluk-107&dil=1. Evidently regional housing expenditure shares are not uniform and vary systematically with income. The correlation between per capita income and housing expenditure share at the NUTS-2 level is 0.73 in 2015, which could be explained by non-homothetic preferences and the discrete, indivisible nature of housing. Abstracting from these channels, we use the median value of 24.5 percent in the model.

costs. In what follows, we refer to the results of this exercise as *short-run* outcomes.

Note that it is possible for some districts to incur welfare losses through trade diversion in the short run. For the parameter values we consider, only two districts—both in Istanbul—experience real wage decreases of 0.8 percent and 1.2 percent. On the side of gains, the median and largest welfare increases are 2.9 percent and 12.4 percent, respectively. As a result, there is substantial heterogeneity in the spatial impact of road improvements in the short-run. However, gains are higher for districts with initially lower real wages. The correlation between these two variables is -0.24. As a result, spatial inequality as captured by the ratio of 90th percentile to median district real wage displays a modest drop from 1.673 to 1.657. Weighted by population, aggregate welfare increase is 2.69 percent in the short-run.

To demonstrate the mechanism through which real incomes are affected in the short run, we calculate for each district a theory-consistent measure of market access:

$$MarketAccess_i = \sum_{j=1}^{N} \hat{T}_{ij}^{1-\sigma} w_j L_j.$$

The scatter plot in Figure B6 visualizes percentage real wage change against percentage change in market access across districts. The correlation between the two variables is 0.97, confirming that locations with larger improvements in market access experienced higher welfare gains in the short run.

Next, we solve for long-run outcomes by holding fixed exogenous parameters $\{\overline{A}_i, \overline{u}_i\}$ and allow population shares $\{L_i\}$ to adjust in response to the improved road network. We do so by solving the system in equation (11) with the national labor market constraint $\sum_i L_i = 1$. Our model predicts an aggregate welfare increase of 2.7 percent in spatial equilibrium when welfare is equalized across districts. Long-run welfare gain is only slightly higher than the population weighted aggregate welfare gain in the short run, which implies that market access rather than the reallocation of labor is the primary driver of the overall impact. When we compare districts' initial population shares with their expected long-run levels, we find that only 1.3 percent of the total population reallocates, consistent with the result that most gains are reaped in the short-run.

Robustness In Table 8, we present the percentage welfare increase in the short- and long-run resulting from travel time reductions for various parameter combinations, highlighting the baseline results in bold. In particular, we report results for when there are no agglomeration economies ($\alpha = 0$) and when agglomeration economies are as strong as permissible ($\alpha = -\beta = 1/3$). We also consider lower and higher elasticities of substitution at $\sigma = 3$ and

$\sigma = 7$. Depending on the parameter combinations, long-run welfare gains vary between 1.8 percent and 5.5 percent.

As expected, welfare increase is larger when differentiated varieties are less substitutable. When goods are highly substitutable, consumers facing high trade costs T_{ij} to other districts incur a lower welfare since they still have access to their local variety at no trade cost, i.e., $T_{ii} = 1$. As α increases, stronger agglomeration economies imply larger welfare gains, although the variation within the permissible range of α values is quite limited. Evidently, the limited amount of long-run labor reallocation that we reported above for baseline outcomes weakens the importance of agglomeration forces. Instead, the elasticity of substitution between varieties plays a preeminent role as the key driver of changes in market access even in the short-run.

Finally, to gauge the role that aggregation plays in our results, we solve the model at higher levels of aggregation, first at the level of 81 provinces and then 26 NUTS-2 regions. For comparability, we keep all parameters at their baseline values except location fundamentals and re-calibrate $(\overline{A}, \overline{u})$ to exactly match provinces' and regions' initial vectors of wages and population shares. To construct the trade cost matrix T_{ij} for inter-provincial trade, we use travel times between central districts of provinces, which are typically urban cores with the highest population. For travel times between regions, we use as nodes the largest provincial center within each NUTS-2 region. The model predicts long-run welfare increases of 2.12 and 1.82 percent at the level of provinces and regions, respectively. These are lower than the 2.7 percent gain when the model was calibrated and solved at the level of districts. As expected from an Armington model of trade with each location producing a distinct variety, increasing the level of aggregation reduces the gains from trade. Having more varieties simply means that improved market access has a bigger role to play.

6.3   Extended Model with Endogenous Labor Supply

In the baseline Allen and Arkolakis (2014) model, location i's population L_i equals its labor supply. The empirical investigation in Section 5 documented increasing employment rates in response to improved market access while there was no such effect on population. To speak to this fact, we now extend the model by allowing endogenous local labor supply to differ from population, and respond to changes in trade costs even when population is immobile in the short-run.[19]

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[19]In a recent paper, Adao, Arkolakis, and Esposito (2021) also highlight the importance of endogenous employment decisions for understanding the effects of aggregate trade shocks on local outcomes in quantitative spatial models. There is also a small literature that introduces endogenous labor supply to standard models of trade, e.g., Neary (1978) and Corsetti, Martin, and Pesenti (2007).

Preferences   The extended model changes household utility from $W_i = Q_i u_i$ in the baseline to an additively separable decision on consumption and labor given by

$$W_i = \left(\frac{Q_i^{1-\frac{1}{\eta}}}{1-\frac{1}{\eta}} - d_i \cdot \frac{h_i^{1+\frac{1}{\gamma}}}{1+\frac{1}{\gamma}} \right) \cdot u_i, \tag{14}$$

where the representative household in location i optimally allocates a fraction h_i of its unit labor endowment to work.[20] This formulation introduces three additional parameters: (γ, η) and location-specific utility cost of labor $\{d_i\}$. The Frisch elasticity γ and η jointly govern the elasticities of labor supply and consumption to the real wage.

The location specific and time-invariant d_i parameter implies spatial heterogeneity in relative taste for consumption versus work. This is needed to fit the spatial variation in employment ratios $\{h_i\}$. We interpret it as a reduced form way of capturing persistent local factors that affect formal labor force participation. These include economic factors, such as commuting costs, and cultural factors such as attitudes toward female labor force participation, age of retirement and self-employment. The literature documents several cases of geographic variation in overall and female labor force participation that could be explained with spatial heterogeneity. Moriconi and Peri (2019) estimate that country-specific labor-leisure preferences explain about a quarter of the variation in employment rates across European countries. Documenting the cross-county dispersion in female labor force participation within the US over time, Fogli and Veldkamp (2011) offer a model of information transmission and preference heterogeneity to explain its dynamics. Such geographic dispersion is potentially relevant for a large developing country like Turkey. To illustrate this, Figure B7 plots the 2006 cross-sectional variation in female employment shares against initial market access. Female employment shares are not only low on average, but they also display a negative spatial correlation with remoteness, leaving room for increased participation in response to market access improvements.

To corroborate this channel, we estimate specification (4) from Section 5 with the change in female employment share as the dependent variable. The results in Table B3 confirm that improved market access is associated with an increase in female employment

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[20]This extension endogenizes labor supply in the participation margin rather than modelling unemployment. A model of unemployment with labor market frictions such as random search and matching would come at a cost of additional complexity whereas our setup allows us to confront labor market outcomes discussed above in a parsimonious way.

share, even when controlling for initial shares in column (2).[21]

Model The household decision can be written as a two stage maximization problem. In the first stage, households chose their location of residence. In the second stage, they make consumption and work decisions, taking local prices and wages as given. For a household residing in i with nominal wage w_i per efficiency unit of labor, the second-stage problem is to maximize the consumption-labor component of the utility function (14) with respect to $\{\{q_{ji}\}_{j=1}^N, h_i\}$ and subject to $\sum_{j=1}^N p_{ji}q_{ji} = w_i h_i$ and $h_i \leq 1$. An interior solution for optimal labor supply in the second-stage is given by

$$h_i = \left(\frac{1}{d_i^\eta} \left(\frac{w_i}{P_i} \right)^{\eta-1} \right)^{\frac{\gamma}{\eta+\gamma}}. \tag{15}$$

Note that for this optimal labor supply response to be consistent with the empirical positive correlation between employment rates and changes in real wages induced by improved market access, the substitution effect between leisure and consumption has to dominate the income effect, i.e., $\eta > 1$.

Substituting equation (15) and the standard CES consumption bundle into the utility function (14) and re-arranging, we get:

$$W_i^* = \frac{1}{\kappa} \cdot d_i^{\frac{\gamma(1-\eta)}{\eta+\gamma}} \cdot \left(\frac{w_i}{P_i} \right)^\kappa \cdot u_i, \tag{16}$$

where

$$\kappa = \frac{(\eta-1)(\gamma+1)}{\eta+\gamma},$$

and P_i is the associated CES price index. In the first stage, households choose location of residence by maximizing W_i^* across all i. Appendix A.1 presents the solution in detail.

The rest of the model and the definition of a spatial equilibrium remains the same as before. Appendix A.2 shows that given trade costs $\{T_{ij}\}$, exogenous components of productivities and amenities $\{\overline{A}_i, \overline{u}_i\}$, disutility of work $\{d_i\}$ and aggregate labor endowment

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[21]This finding is consistent with the results from the literature investigating the effect of trade-induced job opportunities on female labor market participation. Jensen (2012) finds experimental evidence from rural villages in India that job opportunities for women in the business process outsourcing industry led to delayed marriage and childbirth, increased pursuit of education and a higher long-term attachment to the labor market. Heath and Mobarak (2015) similarly find that job growth in the Bangladeshi ready-made garment industry impacted key demographic outcomes for women and increased their labor market participation. Black and Brainerd (2004) finds an association between increased foreign competition and reduced residual gender wage gap across US industries. Expansion of export-oriented manufacturing jobs in Mexico has been associated with increased female labor force participation (Atkin, 2009).

$\overline{L}$, endogenous model outcomes (w_i, h_i, L_i) and level of welfare W, equilibrium constitute a system of $3N+1$ unknowns and non-linear equations. Appendix A.3 proves that the following condition is sufficient for a solution to this system to exist and to be unique:

$$\alpha \cdot \kappa \; + \; \beta \; \leq \; 0, \tag{17}$$

The condition above is reminiscent of the sufficient condition $\alpha + \beta \leq 0$ in the baseline model of Allen and Arkolakis (2014). As $\eta \to \infty$ and $\gamma \to 0$, the extended model converges to the baseline, and so does condition (17) as $\kappa \to 1$. Moreover, the modified sufficient condition has an intuitive interpretation. In this class of models, congestion forces have to be strong enough compared to agglomeration forces for a unique equilibrium to exist. In the baseline model, this is simply a comparison between α and β. In the extended model, agglomeration forces and endogenous labor supply response interact with each other. If values of the two preference parameters (η, γ) governing the responsiveness of labor supply to real wage are such that $\kappa > 1$, then labor supply is highly elastic, and agglomeration forces and labor supply elasticity magnify each other. In that case, the modified condition (17) requires a higher (in absolute value) congestion force β to dominate the combined elasticities of agglomeration and labor supply ($\alpha\kappa$). If, on the other hand, labor supply is less elastic with $\kappa < 1$, then the effect of agglomeration is muted. In that case, the modified condition (17) is relaxed and would be satisfied when congestion force β is weaker or agglomeration force α stronger.

Calibration and Results We set the Frisch elasticity γ equal to 0.4 based on micro-econometric estimates from the extensive margin labor force participation (Reichling and Whalen, 2012; Chetty, Guren, Manoli, and Weber, 2013).

We set η so that the extended model generates the same long-run welfare increase as the baseline model. Inspecting the welfare expression (16), this is the case if the elasticity of welfare to real wage (κ) equals one, as in the baseline utility expression (7). That implies a value of $\eta = 1/\gamma + 2 = 4.5$. Note that keeping (α, β) at their baseline calibration and setting $\kappa = 1$, the sufficient condition (17) for the existence of a unique equilibrium is automatically satisfied. The implied compensated (utility constant) wage elasticity of labor supply from equations (15)-(16) is $(\eta - 1)\gamma/(\eta + \gamma) = 0.36$ which is in line with the estimates of Chetty (2012), and his result that the Frisch elasticity cannot be much larger than the compensated Hicksian elasticity for income effects to be plausible.

Similar to the baseline model, $\{A_i, u_i, d_i\}$ are calibrated by solving the following system of equations using the matrix of trade costs $\hat{T}_{ji}$ estimated by structural gravity and the

observed values of $\{w_i, h_i, L_i\}$:

$$u_i^{\frac{1-\sigma}{\kappa}} \;=\; G_i^{\frac{\sigma-1}{\kappa}} \sum_{j=1}^{N} \hat{T}_{ji}^{1-\sigma} A_j^{\sigma-1} w_i^{\sigma-1} w_j^{1-\sigma}, \tag{18}$$

$$A_i^{1-\sigma} \;=\; \sum_{j=1}^{N} \hat{T}_{ji}^{1-\sigma} L_i^{-1} h_i^{-1} w_i^{-\sigma} w_j^{\sigma} L_j h_j u_j^{\frac{\sigma-1}{\kappa}} G_j^{\frac{\sigma-1}{\kappa}}, \tag{19}$$

$$d_i \;=\; h_i^{-\frac{\eta+\gamma}{\eta\gamma}} \cdot w_i^{\frac{\eta-1}{\eta}} \cdot \left[\sum_{j=1}^{N} \frac{\hat{T}_{ji}^{1-\sigma} w_j^{1-\sigma}}{A_j^{1-\sigma}} \right]^{\frac{1-\eta}{\eta(1-\sigma)}}, \tag{20}$$

where $G_i = (1/\kappa) \cdot d_i^{\frac{\gamma(1-\eta)}{\eta+\gamma}}$. Appendix A.4 presents the derivation of these equations. Note that equation (20) captures the labor supply curve in location i. A high disutility of work d_i shifts it down, resulting in a lower employment ratio h_i at any given wage w_i.

We calibrate location-specific parameters at the level of 81 provinces, which is the intermediate level of aggregation between more than 900 districts and 26 NUTS-2 regions. As discussed in footnote 15, official labor market statistics, including participation, are only available at the NUTS-2 level, which are too large to be functional labor market zones for our purposes. Districts, on the other hand, are typically connected to their provincial centers through commuting ties. Therefore, as geographic labor markets, provinces are the most appropriate level of analysis. Accordingly, we repeat the motivating empirical analysis at the province-level. Table B4 presents the estimation of equation (4) with the change in province-level population and employment ratios as dependent variables, controlling for NUTS-2 level regional fixed effect.[22] The results are consistent with our earlier district-level results in columns (5)-(6) of Table 4 in that population does not respond to reduced travel times but local employment does. The elasticity of province-level employment ratio changes to travel time reductions is -0.288 in column (3) of Table B4.

We finish our analysis by checking whether the calibrated model is capable of generating the reduced-form relationship between reduced travel times and employment ratios documented above. Using the model-implied short-run employment levels at fixed population shares, we replicate the specification from column (3) of Table B4, consistently following the specifications described in footnote 22 above. The elasticity from the model equals -0.083 and is significant at the 5% level, with an R^2 of 0.8. The model-implied elasticity thus cap-

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[22]Similar to the specification in Table 4, we exclude the four provinces (Istanbul, Ankara, Izmir, and Mersin) which already had highway connections in the beginning of the sample period. Since the NUTS-2 region that Mersin belongs to has only two provinces, the other province (Adana) drops from the estimation due to the inclusion of regional fixed effects. As a result, Table B4 has 76 observations.

tures about 30 percent of the empirical elasticity -0.288. The extended model is not only capable of rationalizing its motivating evidence, it is also quantitatively relevant.

7   Conclusion

Developing countries need large investments in transport infrastructure (EBRD, 2017). Yet, evidence on the rates of return for various types of road projects—paving dirt roads, expanding the capacity of existing paved roads, constructing highways—is still scant. We make a contribution to filling this gap by examining the economic benefits of a large-scale public investment program aimed at *expanding the capacity of existing transport infrastructure*. To do so, we look at the case of Turkey – a large upper middle-income country which undertook major public investment in lane-capacity expansion during the 2000s. Our empirical analysis leverages highly disaggregated spatial data on Turkey's road network, high-quality domestic inter-district trade data generated by the universe of domestic firm-to-firm linkages, as well as administrative data on district-level employment and wages.

Our results suggest that travel time reductions due to the ambitious public investment program undertaken by Turkey boosted its intra-national trade, and positively impacted local employment and wages. To gauge the long-run welfare impact, we quantify a workhorse spatial equilibrium model and find aggregate real income gains in the range of 1.8-2.7 percent depending on the level of aggregation. In the short run, when labor is assumed to be immobile, the welfare gains are highly heterogeneous, with the population-weighted aggregate gains amounting to about 2.7 percent.

Our reduced-form empirical findings highlight a novel potential margin of adjustment due to improvements in market access: we find that local employment ratio increased in response to reductions in travel time, with no such effect on population. To rationalize this finding, we extend the model to introduce endogenous labor supply decision. The results imply that the extended model is quantitatively relevant in explaining the reduced-form relationship between reductions in travel times and increasing employment ratios. In particular, the model-implied elasticity captures about one-third of the empirical elasticity.

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Tables and Figures

Table 1: **Summary Statistics**

	2006	2016
Number of source districts	41.8	125.3
	(53.4)	(110.1)
Number of destination districts	51.7	133.8
	(80.2)	(160.3)
Log of trade value (TL)	12.7	12.3
	(1.7)	(2.2)
Period change in log of trade value		1.3
		(1.8)
Period change in log of travel time		-0.14
		(0.06)

Notes: Table shows summary statistics—means and standard errors (in parentheses)—for the number of trade partners, value of and change in bilateral trade between Turkish districts in 2006 and 2016, as well as the change in travel times across them.

Table 2: **Changes in Travel Times and Inter-district Trade**

	(1) $\Delta \ln Trade_{ij}$	(2) $\Delta \ln Trade_{ij}$	(3) $NewTrade_{ij}$	(4) $\Delta Trade_{ij}$
$\Delta \ln TravelTime_{ij}$	-0.736^b	-0.816^b	-0.111^a	-1.614^a
	(0.315)	(0.394)	(0.023)	(0.293)
$\ln Distance_{ij}$	-0.451^a	-0.496^a	-0.0608^a	-0.895^a
	(0.072)	(0.073)	(0.001)	(0.007)
N	20701	15412	555800	555800
R^2	0.341	0.374	0.279	0.289
Origin FE	Y	Y	Y	Y
Destination FE	Y	Y	Y	Y
Province-pair FE	Y	Y	Y	Y

Notes: In column (4), $\Delta Trade$ is calculated using the inverse hyperbolic sine transformation of trade flows. Robust standard errors clustered at the source and destination districts (two-way) are in parentheses. Significance: ^{c} %10, ^{b} 5%, ^{a} 1%.

Table 3: **Changes in Travel Times and Inter-district Trade: Transport Intensity of Industries**

	(1)	(2)	(3)
	$\Delta Trade_{ij}$	$\Delta Trade_{ij}$	$\Delta Trade_{ij}$
$\Delta \ln TravelTime_{ij}$	0.137	0.180	
	(0.654)	(0.649)	
$TimeSensitivity_s * \Delta \ln TravelTime_{ij}$	-1.286	-1.276	-0.793
	(0.806)	(0.802)	(0.958)
$Heavy_s * \Delta \ln TravelTime_{ij}$	-1.549^b	-1.612^b	-1.605^b
	(0.641)	(0.651)	(0.675)
$\ln Distance_{ij}$	-0.305^b	-0.223^c	
	(0.129)	(0.127)	
$TimeSensitivity_s * \ln Distance_{ij}$		0.010	0.066
		(0.132)	(0.142)
$Heavy_s * \ln Distance_{ij}$		-0.127	-0.150
		(0.105)	(0.106)
N	173663	173663	173706
R^2	0.428	0.428	0.284
Origin FE	Y	Y	N
Destination FE	Y	Y	N
Industry-Province pair FE	Y	Y	Y
Origin-Destination FE	N	N	Y

Notes: Sample includes manufacturing industries only, i.e. excludes wholesale. $\Delta Trade$ is calculated using the inverse hyperbolic sine transformation of trade flows. $TimeSensitivity_s$ refers to the share of trade value transported by air in industry s, and $Heavy_s$ captures the (logarithm) of the weight to value ratio. Both measures are constructed by Coşar and Demir (2016) using 2005 UK import data. Robust standard errors clustered at the source and destination districts (two-way) are in parentheses. Significance: ^{c} %10, ^{b} 5%, ^{a} 1%.

Table 4: **Changes in Travel Times and Employment**

	(1)	(2)	(3)	(4)	(5)	(6)
	Total empl.	Manuf. empl.	Manuf. empl.	Manuf. empl.	Population	Ratio of empl. to pop.
$\Delta \ln TravelTime_i^{WgtAvg}$	-2.442^c	-3.379^b			-0.728	-0.670^a
	(1.343)	(1.462)			(0.619)	(0.249)
$\Delta \ln TravelTime_i^{Avg}$			-4.534^b			
			(1.744)			
$\ln Dist_i^{WgtAvg}$	1.155^a	1.333^a	1.389^a	1.183^c	-0.168	-0.138
	(0.385)	(0.492)	(0.501)	(0.623)	(0.253)	(0.098)
$\Delta_{2010-2015} \ln TravelTime_i^{WgtAvg}$				0.585		
				(3.803)		
N	793	715	715	749	793	793
R^2	0.480	0.265	0.268	0.346	0.528	0.428
Province FE	Y	Y	Y	Y	Y	Y

Notes: Table reports results from estimating equation (4). In all columns, except for the last one, dependent variable is the period change in the logarithm of the outcome variable noted in the column title. The period refers to 2006-2016 all columns, except for column (4) where the dependent variable is the change in the logarithm of manufacturing employment between 2006 and 2011. $\Delta \ln TravelTime_i^{WgtAvg}$ and $\Delta \ln TravelTime_i^{Avg}$ are defined in equations (2) and (3). $\Delta_{2010-2015} \ln TravelTime_i^{WgtAvg}$ is constructed using changes in logarithm of bilateral travel times between 2005 and 2010. Except for column (7), the regression is weighted by population size of districts in the beginning of the period (2005). In column (5), initial district-level population share is added as a control variable to equation (4). Large changes (above two log points and below minus one log point) are dropped from the sample. Robust standard errors clustered at the province level are in parentheses. Significance: ^{c} %10, ^{b} 5%, ^{a} 1%.

Table 5: **Changes in Travel Times and Wage per Worker**

	(1)	(2)	(3)
	Total	Manuf.	Manuf.
$\Delta \ln TravelTime_i^{WgtAvg}$	-2.330^a	-1.371^c	
	(0.815)	(0.813)	
$\ln Dist_i^{WgtAvg}$	-0.0741	-0.125	0.446
	(0.172)	(0.282)	(0.268)
$\Delta_{2010-2015} \ln TravelTime_i^{WgtAvg}$			0.0112
			(1.471)
N	795	712	748
R^2	0.326	0.255	0.272
Province FE	Y	Y	Y

Notes: Table reports results from estimating equation (4). In all columns dependent variable is the period change in the logarithm of district-level average wage per worker between 2006-2016 in the first two columns, and between 2006-2011 in the last column. In the first column, wages and employment are aggregated over all industries while they are aggregated over manufacturing and wholesale industries in other columns. $\Delta \ln TravelTime_i^{WgtAvg}$ is defined in equation (2). $\Delta_{2010-2015} \ln TravelTime_i^{WgtAvg}$ is constructed using changes in logarithm of bilateral travel times between 2005 and 2010. The regression is weighted by population size of districts in the beginning of the period (2005). Large changes (above one log point and below minus one log point) are dropped from the sample. Robust standard errors clustered at the province level are in parentheses. Significance: ^{c} %10, ^{b} 5%, ^{a} 1%.

Table 6: **Estimation of Trade Costs**

	(1)	(2)	(3)	(4)
	OLS		WLS	
	2006	2016	2006	2016
$\ln TravelTime_{ij}$	-0.870^a	-1.249^a	-0.776^a	-0.843^a
	(0.018)	(0.043)	(0.046)	(0.053)
N	56,651	164,265	56,651	164,265
R^2	0.597	0.582	0.811	0.796

Notes: All columns include source and destination fixed effects. We set the minimum travel time (including within-districts) to 15 minutes. This is equivalent to assuming that within-district travel times are equal to the minimum travel time across districts. First two columns present OLS estimates and the last two columns present weighted least squares estimates, with trade values used as weights. Robust standard errors clustered (twoway) at the level of source and destination provinces are in parentheses. Significance: ^{c} %10, ^{b} 5%, ^{a} 1%.

Table 7: **Calibration Summary**

Parameter	Explanation	Value	Source/target
α	Agglomeration force	0.07	Average value from Rosenthal and Strange (2004)
β	Congestion force	0.33	Housing share in household expenditures (Turkish Statistical Institute)
σ	Elasticity of substitution	5	Simonovska and Waugh (2014)
T_{ij}	Trade costs	matrix	Structural gravity estimation using travel times
γ	Frisch elasticity of labor supply	0.4	Reichling and Whalen (2012)
η	Preference parameter	4.5	To match baseline welfare response

Table 8: **Long-run Aggregate Welfare Effects**

			σ	
		3	5	7
	0	5.48%	2.70%	1.79%
α	0.07	5.49%	**2.70%**	1.79%
	1/3	5.52%	2.72%	1.81%

Notes: This table reports the aggregate percentage welfare gains for combinations of values for the elasticity of substitution σ and strength of agglomeration economies α. Highlighted cell is the baseline value.

Figure 1: **Turkish Districts and Roads**

Panel A: Road Network in 2005

Panel B: Road Network in 2015

Notes: Data source is Turkish General Directorate of Highways. Black lines denote district boundaries, thin grey lines represent single-carriageway roads, and red lines represent dual-carriageway roads (highways and expressways).

Figure 2: **Turkish Roads over Time**
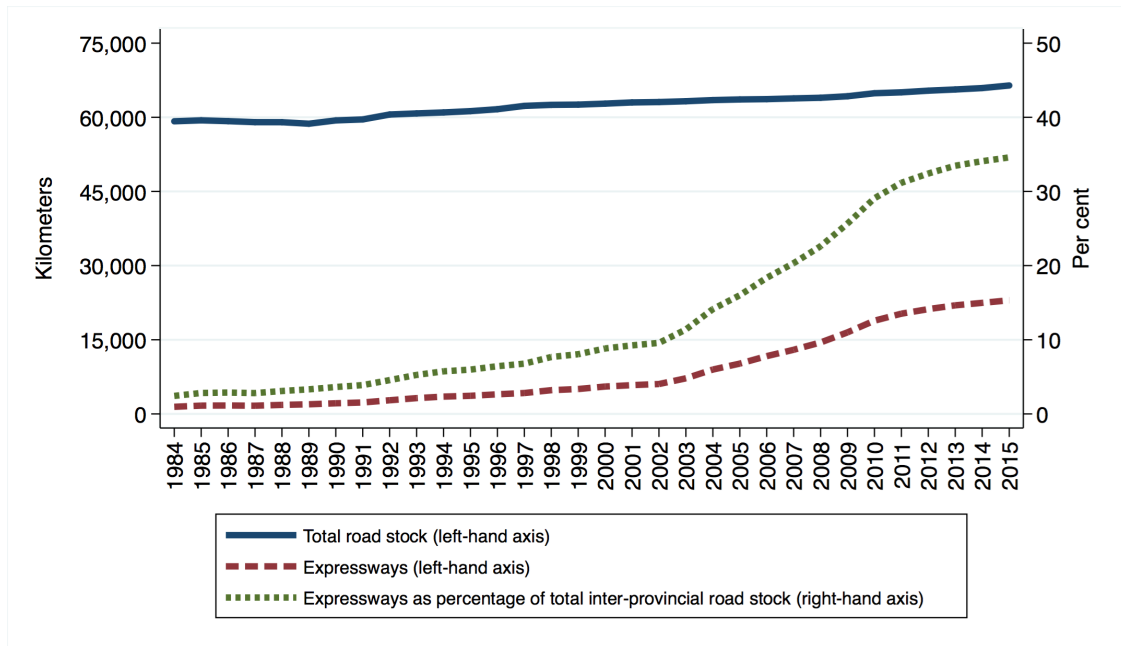
Notes: Data source is Turkish Statistical Institute and General Directorate of Highways. Data downloaded from http://bit.ly/2E3Qh4m, accessed on January 2018.

Figure 3: **Time Savings on Inter-district Travel from 2005 to 2015**

slope=−0.03 R²= 0.09
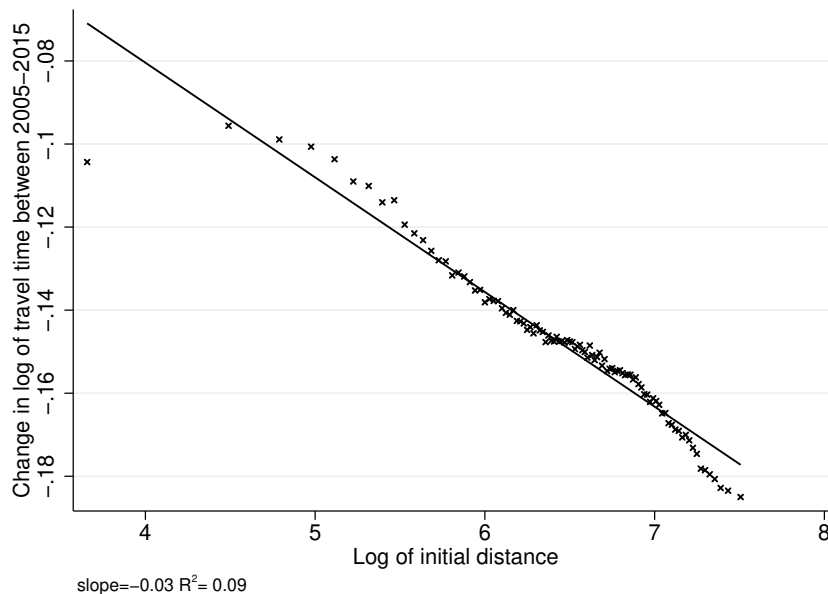
Notes: This chart plots changes in the fastest district-to-district travel times from 2005 to 2015 against the initial distances between them. Each marker represents average distance and change in log of travel time corresponding to a percentile of the distribution of initial bilateral distances between all 913 districts. The reported slope and R^2 are estimated using the full sample.

Figure 4: **Variation in Shortest Travel Times across Districts**

Panel A: Shortest route between Bozkir (Konya) and Ceyhan (Adana)
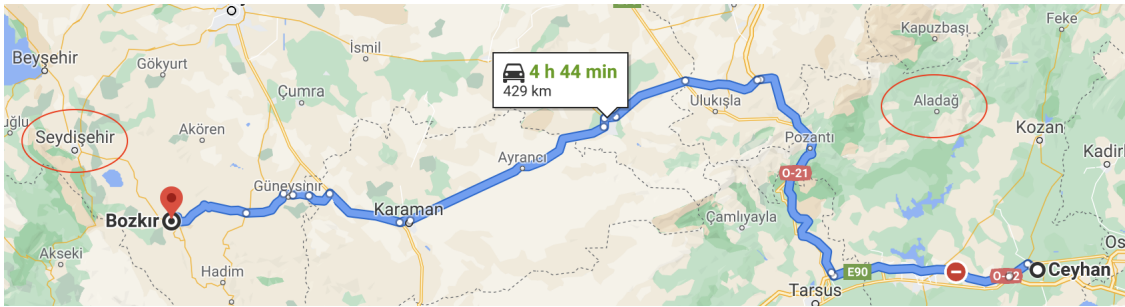
Panel B: Shortest route between Seydisehir (Konya) and Aladag (Adana)
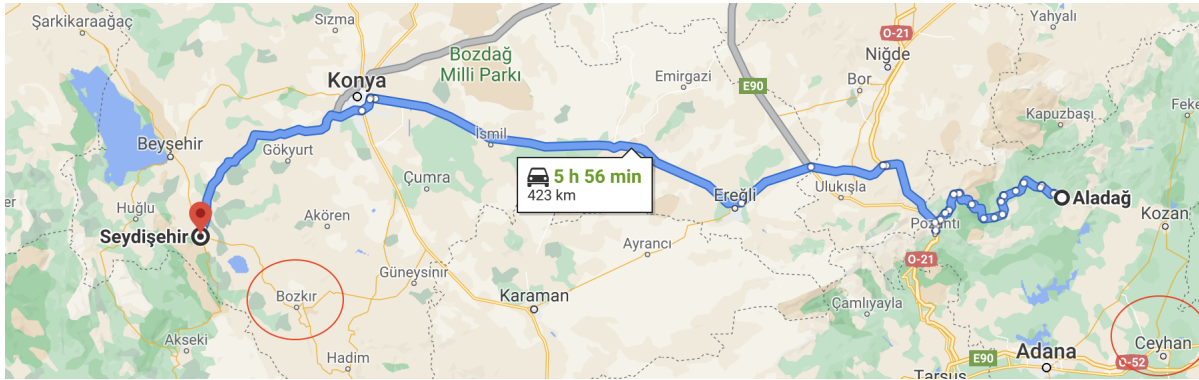
Notes: The upper panel shows the shortest route between Bozkir (Konya) and Ceyhan (Adana), and the lower panel between Seydisehir (Konya) and Aladag (Adana). Source is Google Maps.

Figure 5: **Spatial Distribution of Average Changes in District-level Travel Times**

Panel A: Unadjusted
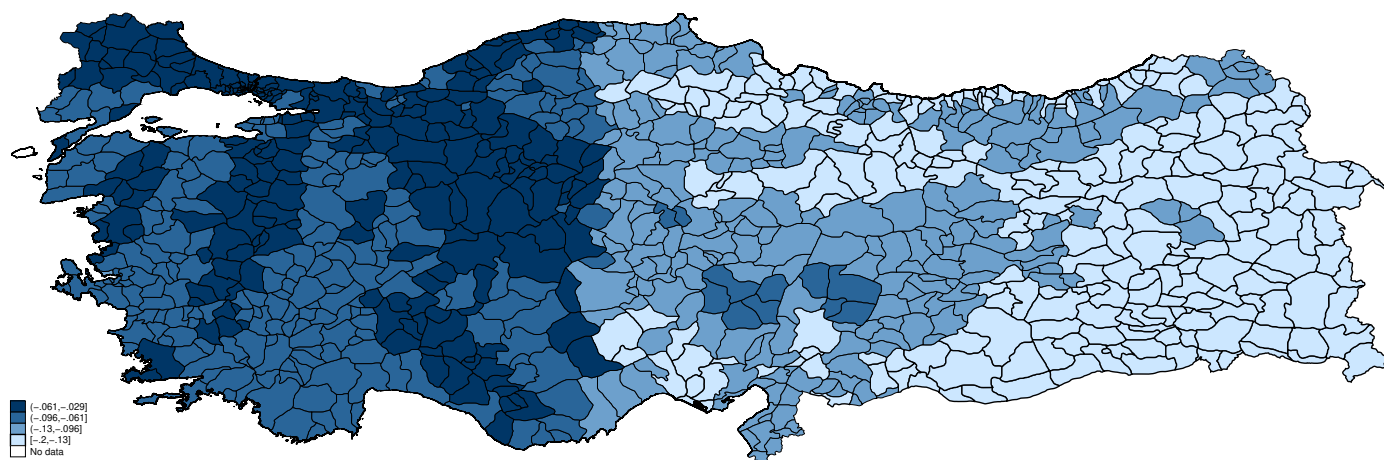
Panel B: Adjusted
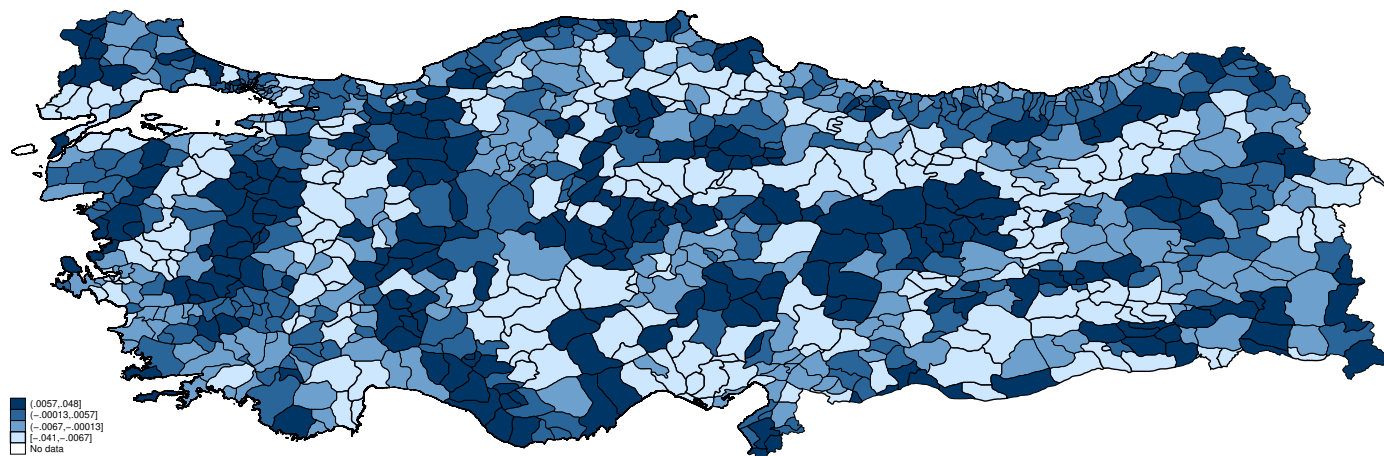
Notes: This figure shows the spatial distribution of the population-weighted average travel time changes at the district level as defined in equation (2). Panel A shows the raw data while Panel B shows the residuals obtained from regressing $\Delta \ln TravelTime_i^{WgtAvg}$ on province fixed effects and population-weighted average of (logarithm of) bilateral distances.

APPENDIX

Equation numbers continue from the main text.

A Extended Model with Labor Supply

A.1 Utility maximization

Consumer problem is to maximize (14) subject to the budget constraint with nominal wage in location i equal to w_i:

$$P_i Q_i = w_i h_i. \tag{21}$$

This can be written as a two-stage maximization problem. First, the consumer maximizes

$$Q_i = \left(\sum_{j=1}^{N} q_{j,i}^{\frac{\sigma-1}{\sigma}} \right)^{\frac{\sigma}{\sigma-1}}$$

subject to $\sum_j p_{j,i} q_{j,i} = P_i Q_i$. Second, we need to determine $P_i Q_i$.

The Lagrangian of the problem is

$$\mathcal{L} = \left(\frac{Q_i^{1-\frac{1}{\eta}}}{1 - \frac{1}{\eta}} - d_i \frac{h_i^{1+\frac{1}{\gamma}}}{1 + \frac{1}{\gamma}} \right) u_i + \lambda(w_i h_i - P_i Q_i)$$

$$\frac{\partial \mathcal{L}}{\partial q_i} = 0 \implies Q_i^{-\frac{1}{\eta}} u_i = \lambda P_i \tag{22}$$

$$\frac{\partial \mathcal{L}}{\partial h_i} = 0 \implies d_i h_i^{\frac{1}{\gamma}} u_i = \lambda w_i \tag{23}$$

Combining (22) and (23), we get:

$$a_i h_i^{\frac{1}{\gamma}} Q_i^{\frac{1}{\eta}} = \frac{w_i}{P_i} \tag{24}$$

Solving for composite consumption yields

$$Q_i = \left(\left(\frac{w_i}{a_i P_i} \right) h_i^{-\frac{1}{\gamma}} \right)^{\eta} \tag{25}$$

Substituting (25) in budget constraint (21),

$$w_i h_i = P_i \left(\left(\frac{w_i}{a_i P_i} \right) h_i^{-\frac{1}{\gamma}} \right)^{\eta}$$

and solving for h_i, we get

$$h_i = \left(\frac{1}{d_i^\eta} \left(\frac{w_i}{P_i} \right)^{\eta-1} \right)^{\frac{\gamma}{\eta+\gamma}} \tag{26}$$

Substituting (26) in (25) yields

$$Q_i = \left(\left(\frac{w_i}{a_i P_i} \right) h_i)^{-\frac{1}{\gamma}} \right)^\eta$$

Substituting for hours worked h_i yields

$$Q_i = d_i^{-\frac{\gamma\eta}{\gamma+\eta}} \left(\frac{w_i}{P_i} \right)^{(\frac{\eta}{\eta+\gamma})(\gamma+1)} \tag{27}$$

Substituting the optimal values of composite consumption (27) and hours worked (26) into utility (14), we get indirect utility as

$$W_i^* = \left(\frac{\eta}{\eta-1} d_i^{\frac{\gamma(1-\eta)}{\gamma+\eta}} \left(\frac{w_i}{P_i} \right)^{\frac{(\eta-1)(\gamma+1)}{\eta+\gamma}} - d_i \frac{\gamma}{\gamma+1} d_i^{-\frac{\eta(\gamma+1)}{\gamma+\eta}} \left(\frac{w_i}{P_i} \right)^{\frac{(\eta-1)(\gamma+1)}{\eta+\gamma}} \right) u_i \tag{28}$$

$$= G_i \left(\frac{w_i}{P_i} \right)^\kappa u_i \tag{29}$$

where

$$G_i = \frac{\eta+\gamma}{(\eta-1)(1+\gamma)} \cdot d_i^{\frac{\gamma(1-\eta)}{\eta+\gamma}} = \frac{1}{\kappa} \cdot d_i^{\frac{\gamma(1-\eta)}{\eta+\gamma}}, \tag{30}$$

which is expression (16) in the main text with $\kappa = \frac{(\eta-1)(\gamma+1)}{\eta+\gamma}$.

A.2   Solving for equilibrium in the extended model

CES demand implies trade flows from j to i equal to

$$X_{ji} = p_{ji} q_{ji} L_i = \left(\frac{\tau_{ji} p_j}{P_i} \right)^{1-\sigma} w_i h_i L_i. \tag{31}$$

Since production is linear in effective labor and competitive in each province, prices are $p_j = w_j / A_j$ at the origin.

Spatial long-run equilibrium holds when wages, population allocation and labor supply decisions $\{w_i, L_i, h_i\}_{i=1}^N$ are such that

- product markets clear:

$$A_i h_i L_i = \sum_{k=1}^N \tau_{ik} q_{ik}, \tag{32}$$

- welfare is equalized across districts:

$$W_i^* = W, \quad \forall i,$$

- aggregate population constraint holds:

$$\sum_{i=1}^{N} L_i = \overline{L}, \tag{33}$$

- districts' expenditures equal their total sales:

$$w_i h_i L_i = \sum_{j=1}^{N} X_{ij}, \tag{34}$$

- households' optimal labor supply is given by:

$$h_i = \left(\frac{1}{d_i^\eta} \left(\frac{w_i}{P_i} \right)^{\eta-1} \right)^{\frac{\gamma}{\eta+\gamma}}. \tag{35}$$

Using equation (31) to substitute out for trade flows and the indirect utility function (16), we can rewrite the market-clearing condition (34) for all $i \in S$ as:

$$
\begin{aligned}
w_i L_i h_i &= \sum_j x_{ij} \\
&= \sum_j \frac{T_{ij} w_i}{A_i P_j}^{1-\sigma} w_j L_j h_j \\
&= \sum_j T_{ij}^{1-\sigma} A_i^{\sigma-1} \left(\frac{w_i}{P_j} \right)^{1-\sigma} \left(\frac{W_j P_j^\kappa}{u_j G_j} \right)^{\frac{1}{\kappa}} L_j h_j \\
L_i h_i w_i^\sigma &= \sum_j T_{ij}^{1-\sigma} A_i^{\sigma-1} \left(\frac{1}{P_j} \right)^{1-\sigma} \left(\frac{W_j P_j^\kappa}{u_j G_j} \right)^{\frac{1}{\kappa}} L_j h_j \\
&= \sum_j T_{ij}^{1-\sigma} A_i^{\sigma-1} \left(\frac{1}{P_j} \right)^{1-\sigma} \left(\frac{W_j^{1-\sigma} W_j^\sigma P_j^\kappa}{u_j G_j} \right)^{\frac{1}{\kappa}} L_j h_j \\
&= \sum_j W_j^{\frac{(1-\sigma)}{\kappa}} T_{ij}^{1-\sigma} A_i^{\sigma-1} P_j^{\sigma-1} \left(\frac{W_j^\sigma P_j^\kappa}{u_j G_j} \right)^{\frac{1}{\kappa}} L_j h_j \\
&= \sum_j W_j^{\frac{(1-\sigma)}{\kappa}} T_{ij}^{1-\sigma} A_i^{\sigma-1} P_j^{\sigma-1} \left(W_j^\sigma \frac{P_j^\kappa}{u_j G_j} \right)^{\frac{1}{\kappa}} L_j h_j \\
L_i h_i w_i^\sigma &= \sum_j W_j^{\frac{(1-\sigma)}{\kappa}} T_{ij}^{1-\sigma} A_i^{\sigma-1} u_j^{\frac{(\sigma-1)}{\kappa}} G_j^{\frac{(\sigma-1)}{\kappa}} L_j h_j w_j^\sigma \tag{36}
\end{aligned}
$$

Combining indirect utility function (16) with the CES price index P_i, we get:

$$w_i^\kappa = \frac{W_i P_i^\kappa}{G_i U_i}$$

$$w_i^{(1-\sigma)\kappa} = W_i^{1-\sigma}P_i^{(1-\sigma)\kappa}G_i^{\sigma-1}U_i^{\sigma-1}$$

$$w_i^{(1-\sigma)\kappa} = W_i^{1-\sigma}G_i^{\sigma-1}U_i^{\sigma-1}\Big(\sum_j T_{ji}^{1-\sigma}A_j^{\sigma-1}w_j^{1-\sigma}ds\Big)^\kappa$$

$$w_i^{(1-\sigma)} = \Big(\sum_j T_{ji}^{1-\sigma}A_j^{\sigma-1}w_j^{1-\sigma}ds\Big)\cdot W_i^{\frac{(1-\sigma)}{\kappa}}G_i^{\frac{(\sigma-1)}{\kappa}}U_i^{\frac{(\sigma-1)}{\kappa}}$$

$$w_i^{(1-\sigma)} = \sum_j W_i^{\frac{(1-\sigma)}{\kappa}}T_{ji}^{1-\sigma}A_j^{\sigma-1}G_i^{\frac{(\sigma-1)}{\kappa}}U_i^{\frac{(\sigma-1)}{\kappa}}w_j^{1-\sigma} \tag{37}$$

When there are productivity and amenity spillovers, and welfare is equalized, substituting equations (6) and (5) into equations (36) and (37) yield:

$$L_i^{1-\alpha(\sigma-1)}h_iw_i^\sigma = W^{\frac{(1-\sigma)}{\kappa}}\sum_j T_{ij}^{1-\sigma}\overline{A}_i^{\sigma-1}\overline{u}_j^{\frac{(\sigma-1)}{\kappa}}G_j^{\frac{(\sigma-1)}{\kappa}}L_j^{1+\beta\frac{(\sigma-1)}{\kappa}}h_jw_j^\sigma \tag{38}$$

$$w_i^{(1-\sigma)}L_i^{\frac{\beta(1-\sigma)}{\kappa}} = W^{\frac{(1-\sigma)}{\kappa}}\sum_j T_{ji}^{1-\sigma}\overline{A}_j^{(\sigma-1)}G_i^{\frac{(\sigma-1)}{\kappa}}\overline{u}_i^{\frac{(\sigma-1)}{\kappa}}w_j^{1-\sigma}L_j^{\alpha(\sigma-1)}ds \tag{39}$$

Using the CES price index P_i, (33), (15), (38), and (39), we can solve for L_i, w_i, h_i, P_i and W.

A.3 Proof of existence and uniqueness of equilibrium

Combining (33) and (15) yields:

$$h_i = \left(\frac{1}{d_i^\eta}\left(\frac{W}{u_iG_i}\right)^{\frac{\eta-1}{\kappa}}\right)^{\frac{\gamma}{\eta+\gamma}} \tag{40}$$

Substituting equation (40) into (43) and solving for w_i yields:

$$w_i = \phi^{\frac{1}{2\sigma-1}}L_i^{\frac{1}{1-2\sigma}}A_i^{\tilde{\sigma}}d_i^{c_1}W^{c_2}(u_iG_i)^{\frac{c_2}{2\sigma-1}-\frac{\tilde{\sigma}}{\kappa}} \tag{41}$$

where $c_1 = \frac{-\eta\gamma}{(\eta+\gamma)(1-2\sigma)}$, $c_2 \equiv \frac{\eta-1}{\kappa}\frac{\gamma}{\eta+\gamma}\frac{1}{1-2\sigma}$ and $\tilde{\sigma} \equiv \frac{\sigma-1}{2\sigma-1}$. Note that, in the proofs before we have used the notion of discrete spaces, but all the above relationships hold for continuous spaces as well, so in all the above equations sum can be interchangeably replaced by integrals. For this proof, we will refer to the properties of nonlinear integral equations. Now, substituting equations (41), (5) and (6) into (38) (or (39)) yields:

$$L_i^{\tilde{\sigma}\gamma_1}\bar{A}_i^{\tilde{\sigma}(1-\sigma)}d_i^{c_1(1-\sigma)}(\bar{u}_iG_i)^{-c_2\tilde{\sigma}-\frac{\tilde{\sigma}(1-\sigma)}{\kappa}}$$
$$= W^{\frac{(1-\sigma)}{\kappa}}G_i^{\frac{(\sigma-1)}{\kappa}}\bar{u}_i^{\frac{(\sigma-1)}{\kappa}}\int_j T_{ji}^{1-\sigma}\bar{A}_j^{\sigma-1}L_j^{\tilde{\sigma}\gamma_2}\bar{A}_j^{\tilde{\sigma}(1-\sigma)}a_s^{c_1(1-\sigma)}(\bar{u}_jG_s)^{-c_2\tilde{\sigma}-\frac{\tilde{\sigma}(1-\sigma)}{\kappa}}ds$$

where

$$\gamma_1 \equiv 1 - \alpha(\sigma-1) - \frac{\beta\gamma}{(\gamma+1)(1-2\sigma)} - \frac{\beta}{\kappa}\sigma,$$

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$$\gamma_2 \equiv 1 + \alpha\sigma - \frac{\beta\gamma}{(\gamma+1)(1-2\sigma)} + (\sigma-1)\frac{\beta}{\kappa}.$$

Equivalently,

$$
\begin{aligned}
L_i^{\tilde{\sigma}\gamma_1} &= \bar{A}_i^{\tilde{\sigma}(\sigma-1)} d_i^{c_1(\sigma-1)} \left(\bar{u}_i G_i\right)^{c_2\tilde{\sigma}+\frac{(1-\tilde{\sigma})(\sigma-1)}{\kappa}} W^{\frac{(1-\sigma)}{\kappa}} \\
&\times \int_j T_{ji}^{1-\sigma} \bar{A}_j^{(\sigma-1)(1-\tilde{\sigma})} d_s^{c_1(1-\sigma)} \left(\bar{u}_j G_s\right)^{\frac{\tilde{\sigma}(\sigma-1)}{\kappa}-c_2\tilde{\sigma}} \left(L_j^{\tilde{\sigma}\gamma_1}\right)^{\frac{\gamma_2}{\gamma_1}} ds. \quad (42)
\end{aligned}
$$

This equation resembles equation (13) of Allen and Arkolakis (2014) when $\gamma = 0$ and $\eta \to \infty$.

A.3.1 Existence

We can rewrite equation (42) as a nonlinear integral equation

$$f_i = \lambda \int_j K_{ji} f_j^{\frac{\gamma_2}{\gamma_1}} ds,$$

where $f_i \equiv L_i^{\tilde{\sigma}\gamma_1}$, $\lambda \equiv W^{\frac{(1-\sigma)}{\kappa}}$ and

$$K_{ji} \equiv \bar{A}_i^{\tilde{\sigma}(\sigma-1)} d_i^{c_1(\sigma-1)} \left(\bar{u}_i G_i\right)^{c_2\tilde{\sigma}+\frac{(1-\tilde{\sigma})(\sigma-1)}{\kappa}} T_{ji}^{1-\sigma} \bar{A}_j^{(\sigma-1)(1-\tilde{\sigma})} d_j^{c_1(1-\sigma)} \left(\bar{u}_j G_j\right)^{\frac{\tilde{\sigma}(\sigma-1)}{\kappa}-c_2\tilde{\sigma}}.$$

Then, the arguments in AA hold here as well once we state assumptions on preference weight d_i, where d_i is continuous and bounded above and below by strictly positive numbers. Recall that in their Theorem 2, Allen and Arkolakis (2014) consider a regular geography, where $\bar{A}$, $\bar{u}$ and T are continuous and bounded above and below by strictly positive numbers. Thus, K_{ji} is bounded above and below by a positive number, so Theorem 2 of Karlin and Nirenberg (1967) also applies to our setting, completing the proof of existence.

A.3.2 Uniqueness

The only difference from AA is in K_{ji}, but under additional assumption on d_i every argument in proving uniqueness holds in our setup as well. That is, if $\left| \frac{\gamma_2}{\gamma_1} \right| \leq 1$ then uniqueness is guaranteed. Next, we characterize the parameter space in terms of α, β, σ, γ and η.

Rewrite

$$\frac{\gamma_2}{\gamma_1} = \frac{1 - \lambda + \alpha\sigma + (\sigma-1)\frac{\beta}{\kappa}}{1 - \lambda - \alpha(\sigma-1) - \sigma\frac{\beta}{\kappa}},$$

where $\kappa = \frac{(\eta-1)(1+\gamma)}{\eta+\gamma}$ and $\lambda = \frac{\beta\gamma}{(\gamma+1)(1-2\sigma)}$.

One can show that $\left| \frac{\gamma_2}{\gamma_1} \right| \leq 1$ if and only if

$$\left(\frac{\beta}{\kappa} + \alpha < \frac{1-\lambda+\alpha}{\sigma} \text{ and } \frac{\beta}{\kappa} + \alpha \leq \min\{2(1-\lambda+\alpha), 0\} \right) \text{ if } \gamma_1 > 0$$

$$\left(\frac{\beta}{\kappa} + \alpha > \frac{1-\lambda+\alpha}{\sigma} \text{ and } \frac{\beta}{\kappa} + \alpha \geq \max\{2(1-\lambda+\alpha), 0\} \right) \text{ if } \gamma_1 < 0.$$

To be consistent with Allen and Arkolakis (2014), consider the case where $\gamma_1 > 0$. Since $-1 < \beta < 0$, the elasticity of hours worked with respect to real wages $\gamma > 0$, $\sigma > 1$, $0 \leq \lambda \leq 1$. Since $\alpha > 0$, $1 - \lambda + \alpha > 0$, the condition for uniqueness in this case boils down to $\frac{\beta}{\kappa} + \alpha \leq 0$.

A.4 Solving for exogenous amenities, productivities, and preferences for labor force participation

Now, we derive the equations required to solve for the exogenous amenities and productivities. To do this, similar to Allen and Arkolakis (2014) we first can show that if (43) below holds and transportation costs are symmetric, then if either (38) or (39) holds, the other one holds as well.

$$L_i h_i w_i^\sigma A_i^{1-\sigma} = \Phi w_i^{1-\sigma} (u_i G_i)^{\frac{1-\sigma}{\kappa}} \tag{43}$$

Combined with welfare equalization across regions, we can rewrite (37) as:

$$u_i^{\frac{1-\sigma}{\kappa}} = W^{\frac{1-\sigma}{\kappa}} G_i^{\frac{\sigma-1}{\kappa}} \sum_j T_{ji}^{1-\sigma} A_j^{\sigma-1} w_i^{\sigma-1} w_j^{1-\sigma} ds \tag{44}$$

This is the equation to solve for amenity.

Combining the CES price index with (43),

$$
\begin{aligned}
A_i^{1-\sigma} &= \Phi w_i^{1-\sigma} (u_i G_i)^{\frac{1-\sigma}{\kappa}} L_i^{-1} h_i^{-1} w_i^{-\sigma} \\
&= \Phi W^{\frac{1-\sigma}{\kappa}} \sum_j T_{ji}^{1-\sigma} (A_j^{\sigma-1} w_j^{1-\sigma}) L_i^{-1} h_i^{-1} w_i^{-\sigma} ds
\end{aligned}
$$

Using (43) to replace $(A_j^{\sigma-1} w_j^{1-\sigma})$ we get

$$A_i^{1-\sigma} = W^{\frac{1-\sigma}{\kappa}} \sum_j T_{ji}^{1-\sigma} L_i^{-1} h_i^{-1} w_i^{-\sigma} w_j^\sigma L_j h_j (u_j G_j)^{\frac{\sigma-1}{\kappa}} ds \tag{45}$$

This is the equation to solve for productivity. Combining (30), (15), and the CES price index, and after simplifying, we get the the following equation to solve for d_i:

$$d_i = h_i^{-\frac{\eta+\gamma}{\eta\gamma}} \cdot w_i^{\frac{\eta-1}{\eta}} \cdot \left[\sum_{j=1}^N \left(\frac{\hat{T}_{ji} w_j}{A_j} \right)^{1-\sigma} \right]^{\frac{1-\eta}{\eta(1-\sigma)}} \tag{46}$$

B   Additional Tables and Figures

Table B1: **Changes in Travel Times and Inter-district Trade: Robustness Checks**

	(1)	(2)	(3)
	$\Delta \ln Trade_{ij}$	$\Delta \ln Trade_{ij}$	$\Delta Trade_{ij}$
$\Delta \ln TravelTime_{ij}$	-0.243	-1.675[a]	-1.764[a]
	(0.958)	(0.492)	(0.286)
$\ln Distance_{ij}$	-0.558[a]	-0.672[a]	-0.933[a]
	(0.147)	(0.091)	(0.012)
N	11628	6679	549287
R^2	0.354	0.440	0.236
Origin FE	Y	Y	Y
Destination FE	Y	Y	Y
Province-pair FE	Y	Y	Y

Notes: In column (1), dependent variable is calculated over the 2006-2011 period while change in travel times between districts are calculated over the 2011-2016 period. Columns (2) and (3) exclude multi-establishment firms from either side of the transaction before aggregating trade flows at the district level. In column (3), $\Delta Trade$ is calculated using the inverse hyperbolic sine transformation of trade flows. Robust standard errors clustered at the source and destination districts (two-way) are in parentheses. Significance: [c] %10, [c] 5%, [a] 1%.

Table B2: **Changes in Travel Times and Initial District-level Characteristics**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	$\Delta \ln TravelTime_i^{WgtAvg}$			$\Delta \ln TravelTime_i^{WgtAvg}$			$\Delta \ln TravelTime_i^{Avg}$
Population Density$_{i,t=2005}$	0.0046[a]	0.0018	0.0002			0.0002	0.0002
	(0.0017)	(0.0013)	(0.0007)			(0.0003)	(0.0003)
Ratio of employment	0.106[a]	0.0364[c]		-0.0162		-0.0181	-0.0145
to population$_{i,t=2006}$	(0.0285)	(0.0192)		(0.0111)		(0.0111)	(0.0106)
$\ln Dist_i^{WgtAvg}$		-0.0636[a]			-0.0119	-0.0171	-0.0032
		(0.00907)			(0.0182)	(0.0175)	(0.0157)
N	793	793	793	793	793	793	793
R^2	0.150	0.391	0.832	0.833	0.832	0.833	0.801
Province FE	N	N	Y	Y	Y	Y	Y

Notes: In all columns, except for the last one, dependent variable is the weighted average of the logarithmic change in bilateral travel times between 2005-2015 as defined in equation (2). In the last column, it is the unweighted average as defined in equation (3). Population density is defined as 1,000 people per square km. Robust standard errors clustered at the province level are in parentheses. Significance: [c] %10, [b] 5%, [a] 1%.

Table B3: **Changes in Travel Times and Female Employment Share**

	(1)	(2)
$\Delta \ln TravelTime_i^{WgtAvg}$	-0.291[b]	-0.404[a]
	(0.139)	(0.127)
$\ln Dist_i^{WgtAvg}$	-0.108[a]	-0.161[a]
	(0.0362)	(0.0425)
Initial female share$_i$		-0.329[a]
		(0.0354)
N	795	795
R^2	0.237	0.390
Province FE	Y	Y

Notes: Table reports results from estimating equation (4). In all columns dependent variable is the period change in the share of female employment in total district-level employment between 2006-2016. $\Delta \ln TravelTime_i^{WgtAvg}$ is defined in equation (2). The regression is weighted by population size of districts in the beginning of the period (2005). Robust standard errors clustered at the province level are in parentheses. Significance: [c] %10, [b] 5%, [a] 1%.

Table B4: **Changes in Travel Times and Employment: Province-level Results**

	(1)	(2)	(3)
	Employment	Population	Ratio of empl. to pop.
$\Delta \ln TravelTime_i^{WgtAvg}$	-4.011[c]	-1.800	-0.288[c]
	(2.021)	(1.511)	(0.155)
$\ln Dist_i^{WgtAvg}$	0.102	0.191	-0.0345
	(0.436)	(0.142)	(0.0524)
N	76	76	76
R^2	0.695	0.612	0.472
NUTS2 FE	Y	Y	Y

Notes: Table reports results from estimating equation (4) at the level of provinces. In all columns, except for the last one, dependent variable is the period change in the logarithm of the outcome variable noted in the column title. The period refers to 2006-2016 all columns. $\Delta \ln TravelTime_i^{WgtAvg}$ measures the change in average travel time between a province and all other 80 provinces of the country between 2005-2015, weighted by the initial population of destination provinces. Except for column (2), the regression is weighted by population size of provinces in the beginning of the period (2005). Robust standard errors clustered at the NUTS2 level are in parentheses. Significance: [c] %10, [b] 5%, [a] 1%.

Figure B1: **Employment in Private Sector across Provinces**

Panel A: 2011
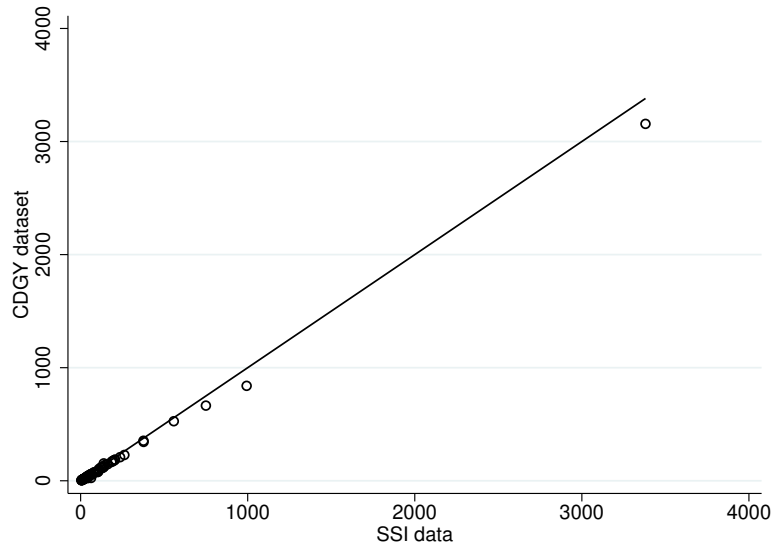
Panel B: 2016
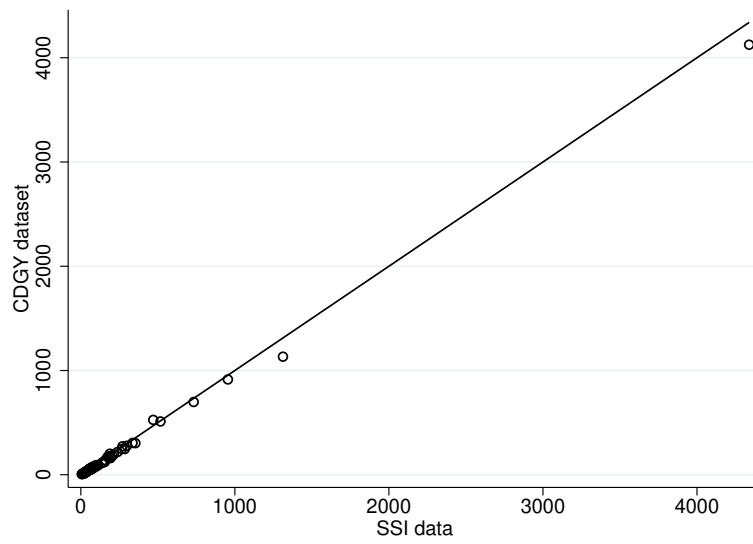
Notes: These figures compare the province-level formal employment in private sector calculated from our data to the publicly available data reported by SSI. Panel A presents the comparison for the year 2011, and Panel B for 2016. All numbers are in 1,000 workers.

Figure B2: **Distribution of $\hat{\theta}_1$**

Panel A: Intensive margin
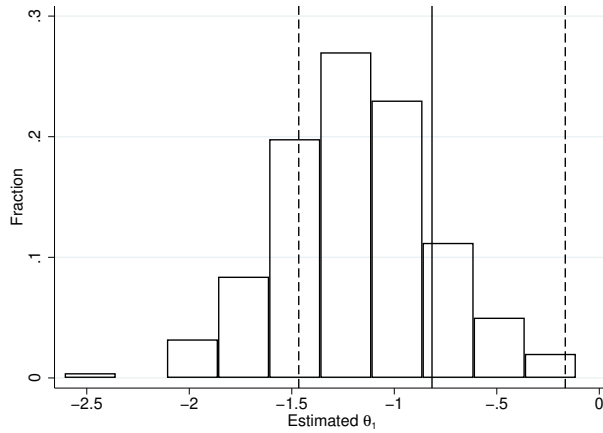
Panel B: Intensive and extensive margins
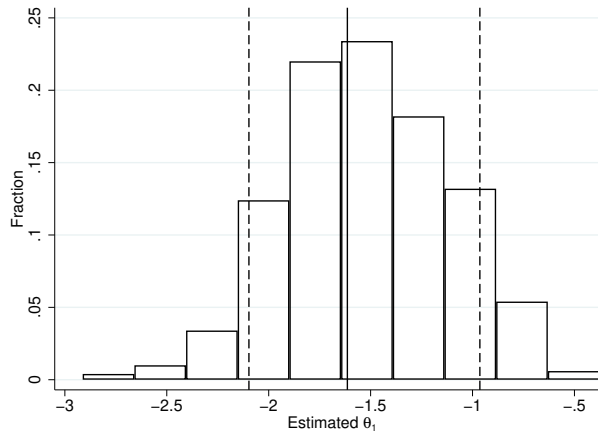
Notes: This figure plots the distribution of the estimate of θ_1 in equation (1), obtained from estimating the equation on 500 randomly drawn samples of district pairs, with sample size equal to one-third of the baseline sample. The empirical specification presented in Panel A corresponds to column 2 of Table 2, and Panel B corresponds to column 4. Dashed lines in both panels correspond to the lower and upper bounds of the 90% confidence interval for the respective specification in Table 2. Similarly, the solid lines represent the respective point estimates.

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Figure B3: **Changes in Trade Flows between Turkish Districts (2006-2016)**

Panel A: Predicted and Actual Changes in Trade Flows
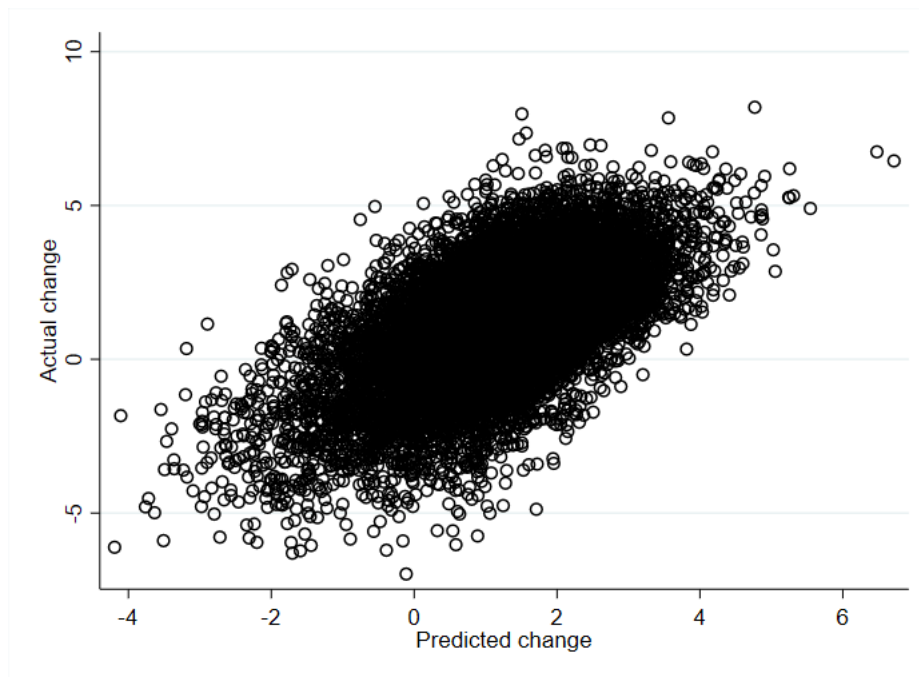
Panel B: Changes in Trade Flows and Travel Times
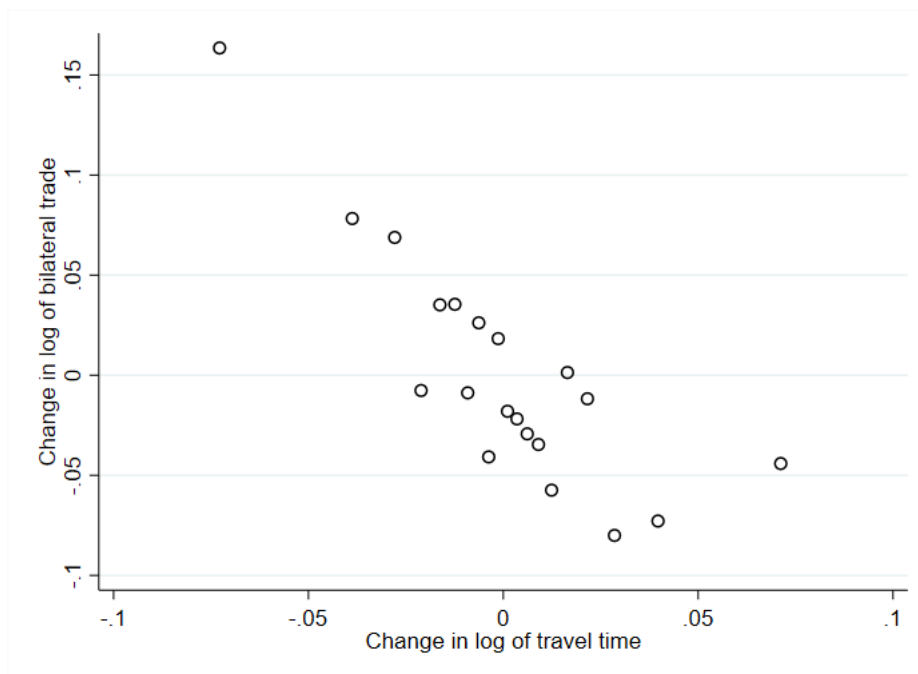
Notes: The figure in the upper panel presents the actual and predicted period change in district-level bilateral domestic trade flows. The predicted values are based on the estimates presented in column (2) of Table 2. The lower panel presents a scatter plot of the vigintiles of the period changes in the logarithm of bilateral trade flows against changes in bilateral travel times. Both series are residualized with respect to source, destination, and province-pair fixed effects, as well as the logarithm of bilateral distance.

Figure B4: **Changes in Travel Times under Alternative Speed Assumptions**

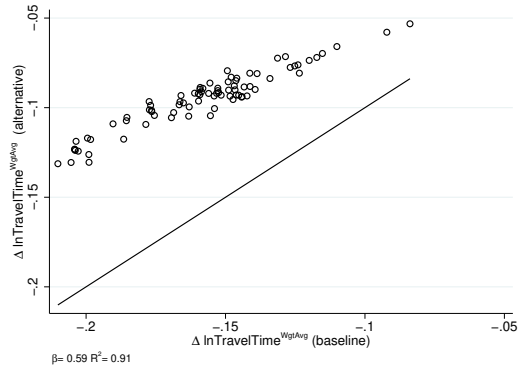(a) Changes in bilateral travel times (Scenario 1)
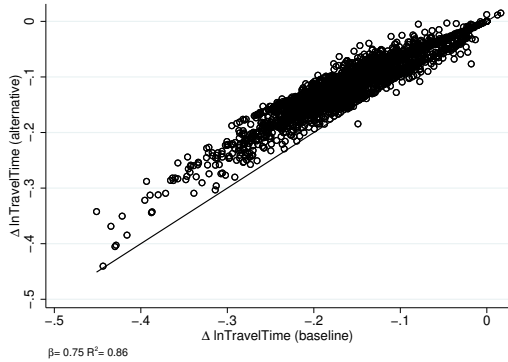
(b) Changes in bilateral travel times (Scenario 2)
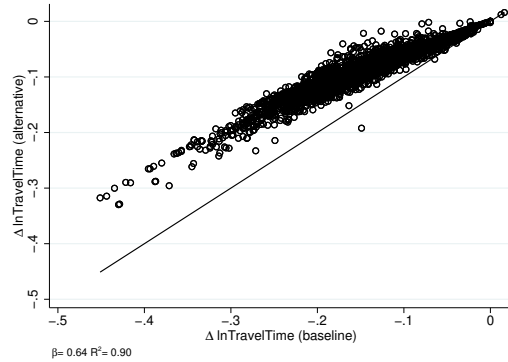
(c) Changes in population-weighted travel times (Scenario 1)
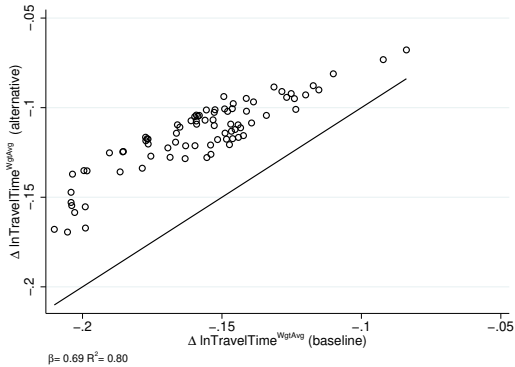
(d) Changes in population-weighted travel times (Scenario 2)

Notes: These figures show period changes in bilateral travel times between Turkish provinces (panels (a) and (b)) and population-weighted changes in travel times at the province level (panels (c) and (d)) under two alternative sets of assumptions about average speeds on the following three types of roads: highways, four-lane expressways, and single carriageways. Under the baseline scenario, average speed per hour is assumed to be 100 km on highways, 90 km on four-lane expressways, and 65 km on single carriageways. Under the first alternative scenario, these speed assumptions are 110 km, 80 km, and 65 km; and under the second scenario, they are 110 km, 90 km, and 75 km.

Figure B5: **Calibrated Exogenous Characteristics of Districts**
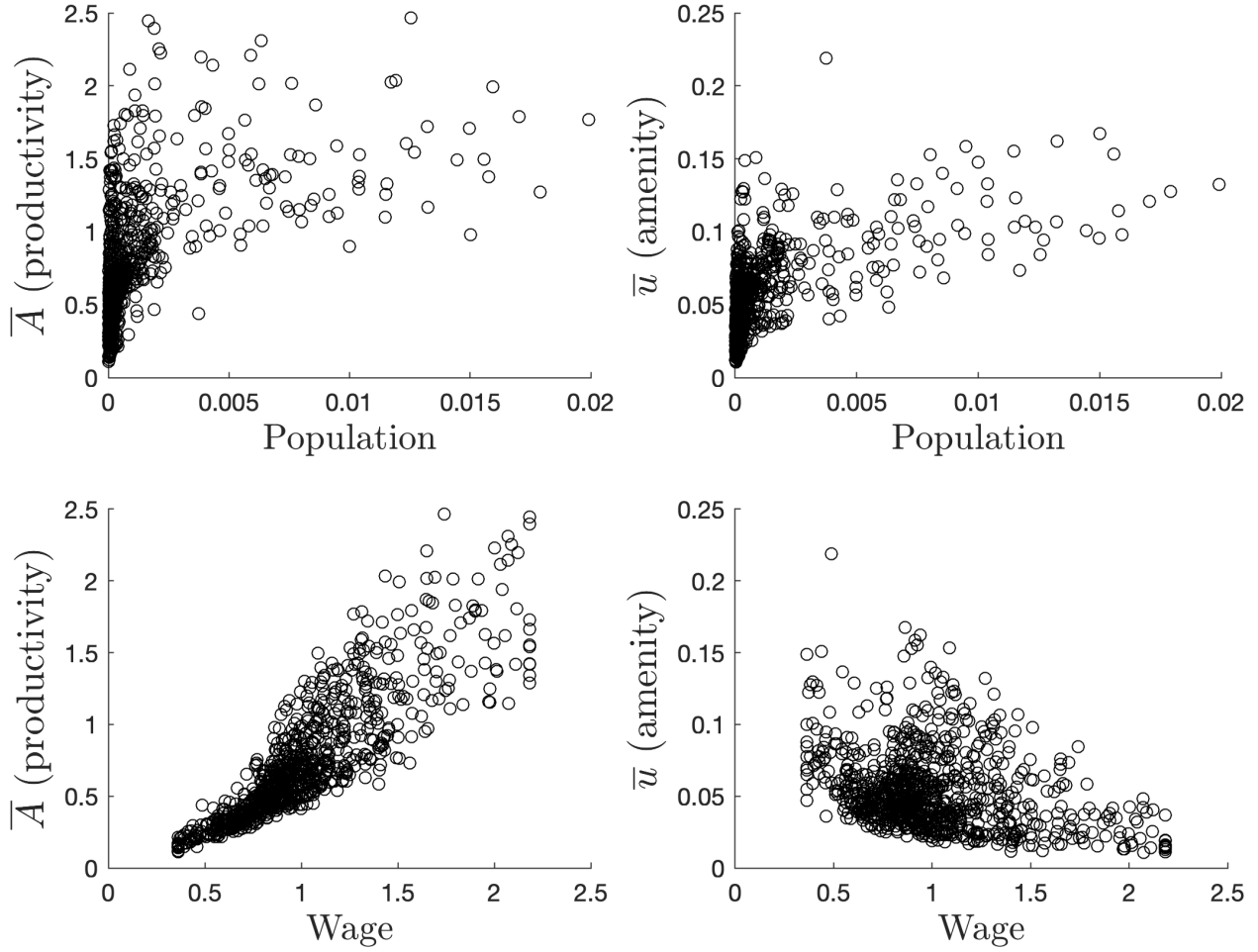
Notes: Each observation is a district. Population and wages are districts' urban population shares and normalized per worker wages in 2006. $\overline{A}$, and $\overline{u}$ are the exogenous productivities and amenities, respectively.

Figure B6: **Short-run Changes in Market Access and Real Wage**
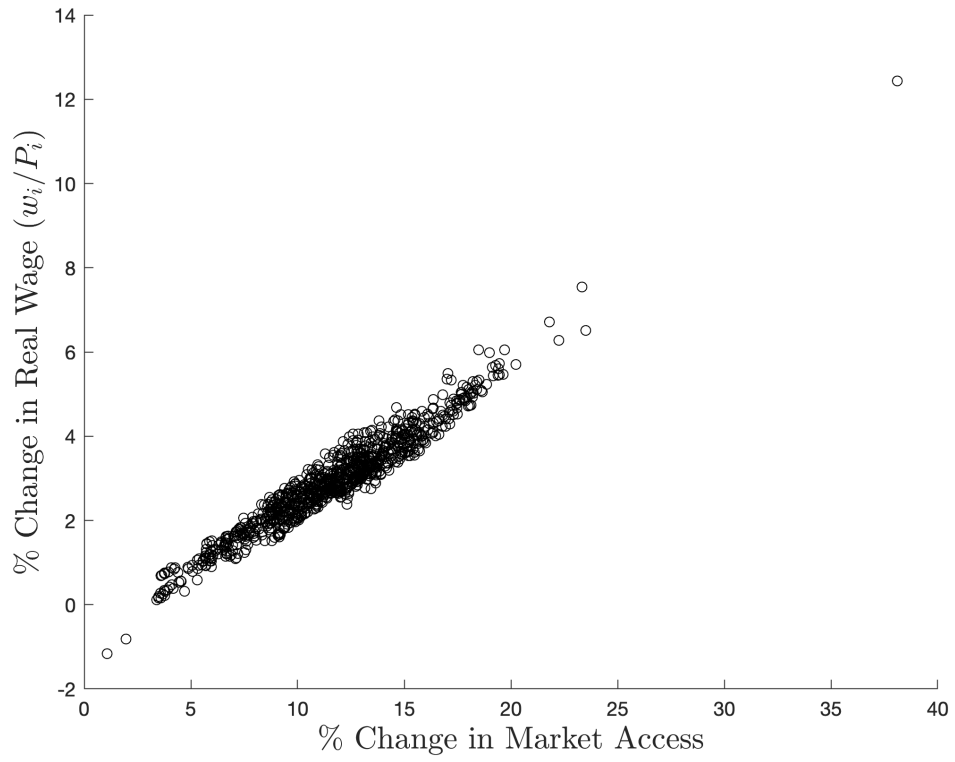
Notes: Each observation is a district. The y-axis is the percentage change in real wage (w/p) when labor is immobile in the short-run. The x-axis is the percentage change in market access defined as $MarketAccess_i = \sum_{j=1}^{N} \hat{T}_{ij}^{1-\sigma} w_j L_j$. See Section 6.2 for details.

Figure B7: **Female Employment Share and Initial Market Access**

Panel A: All districts
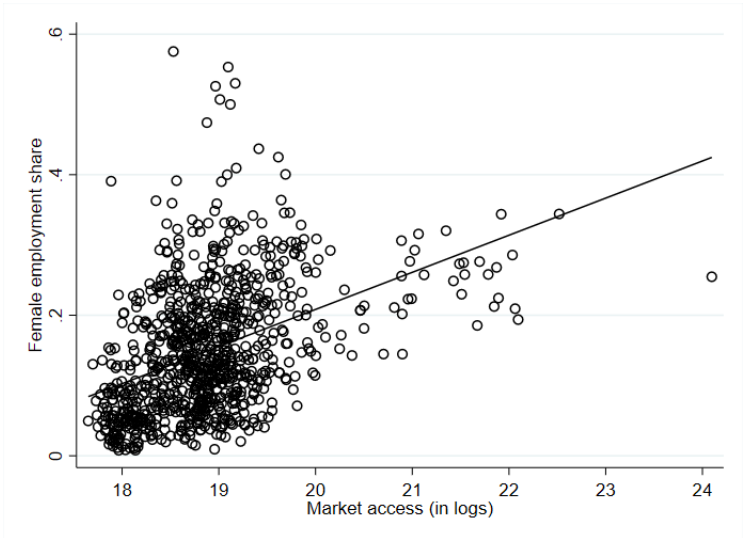
Panel B: Excluding major port cities
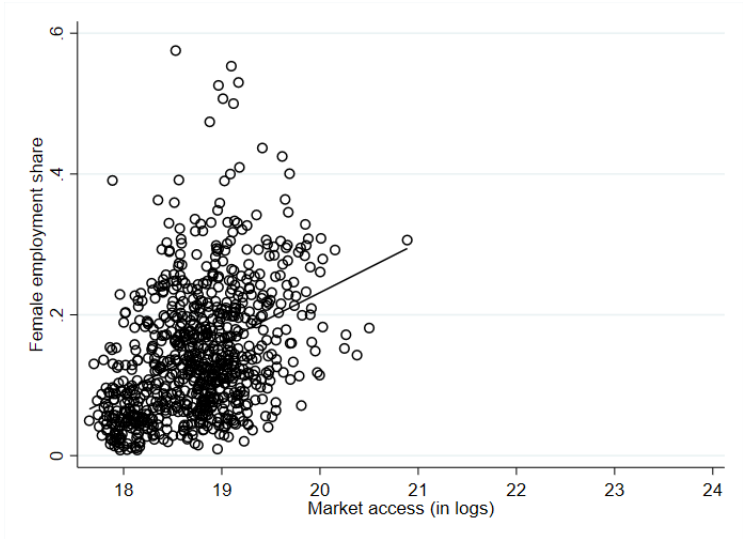
Notes: These figures show the female share of employment at the district level against the initial market access defined as $\sum_j (w_j^{2005} L_j^{2005})/T_{ij,2005}^{-1}$. The estimated linear slope is 3.36 (s.e. 0.24) in the upper panel and 2.27 (s.e. 0.18) in the lower panel.