# GLOBALIZATION, AGRICULTURAL MARKETS AND MASS MIGRATION: ITALY, 1881-1912

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**ABSTRACT**

Despite the significant attention paid to the current consequences of globalization for migration behavior, there are few historical accounts of the effect of commodity market integration at the local level. We set our paper within the context of the first globalization era, when migration flows were largely unregulated, and highlight how exogenous shocks in agricultural commodity prices influenced international migration flows from Italian provinces between 1881 and 1912. To do this, we construct an index of global price exposure based on the initial provincial agricultural production structures. Our analysis quantifies the contribution of globalization-induced agricultural-price shocks to migration decisions, alongside more traditional explanatory factors such as migrant networks and landholding systems. We find evidence that agricultural-price shocks are positively related to the propensity to migrate, as migration tended to increase in proportion with agricultural commodity prices. This result suggests that liquidity constraints were binding until agricultural incomes reached a certain threshold. These findings can inform our understanding of present-day migration responses in developing countries in the face of even more rapid globalization but higher barriers to legal migration.

**Key Words**: Age of mass migration; determinants of migration; agricultural-price shocks
**JEL Codes**: N93; N13; F22; O15

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## 1. INTRODUCTION

The consequences of globalization for migration have been recently explored in the developing country context. For example, Dix-Carneiro and Kovak (2015) find that Brazilian trade liberalization had little effect on migration, while Majlesi and Narciso (2018) provide evidence that Chinese import competition reduced Mexican migration to the US. Our paper takes an historical perspective and investigates the effect of globalization-induced agricultural-price shocks on international out-migration from Italian provinces over the period 1881 to 1912. This period was free of legal barriers to migration, allowing us to focus on the theoretical tension between two opposite mechanisms through which price shocks might affect migration. First, negative agricultural-price changes may reduce agricultural incomes and create a labor surplus that may not be absorbed in the infant industrial sector, thus pushing people towards migration. Second, negative agricultural-price changes may reduce agricultural incomes and limit the ability to cover the costs of migration.

Our analysis provides evidence that agricultural-price shocks are positively related to the propensity to migrate and suggests that the liquidity constraints channel dominated in the historical Italian case. We interpret liquidity constraints in line with the argument by Faini and Venturini (1994) that Italy was caught in a poverty trap that curbed out-migration, because people could not finance their moves and relied heavily on remittances and prepaid tickets to achieve a move. We thus provide new evidence on the importance of liquidity constraints during this period, which have been previously postulated to explain the puzzle identified by Hatton and Williamson (2005). They showed the general pattern that the initial immigrant source countries during the first global era were all in northwest Europe, which had lower wage differentials with the US than southern countries like Italy. Furthermore, migration rates typically increased as economic growth occurred in the source countries, suggesting that liquidity constraints must be large enough to counteract closing wage gaps.[1] Papers such as Covarrubias et al (2015) have found some evidence of liquidity constraints using annual, country-level migration flows and aggregate GDP data, but in this paper we present evidence based on finer data and a more plausibly exogenous shock to incomes.

[1] Hatton (2010) highlights the difference between countries like Ireland, where narrowing wage differentials reduced emigration, and Italy, where convergence was lower and also the liquidity trap kept emigration high. Using modern data, Belot and Hatton (2012) show that any study of migrant selection must take into account the fact that those with the least skill typically cannot afford to migrate.

In order to estimate whether agricultural-price shocks foster or restrain migration within the above-outlined framework, we construct an index of international price exposure – a proxy for commodity market integration during the first globalization era[2] – based on the agricultural production structure of each province in Italy. The intuition behind this measure is that international agricultural-price shocks affect each geographical unit of analysis differently, on the basis of the share of land allocated to each specific crop. Our analysis suggests that migration increased proportionally to agricultural prices. For example, looking at a period of rising prices, from 1893 to 1900, we estimate that the resulting boost to agricultural incomes can explain up to one half of the rise in the total migration rate at that time. We argue that higher agricultural commodity prices translated into higher income, relaxing liquidity constraints and allowing would-be migrants to afford passage abroad.[3]

We also investigate the role of income uncertainty – proxied by agricultural commodity price fluctuations – and show how increased uncertainty tended to increase migration rates. Overall, our results suggest that Italians were more sensitive to changes in the level of income than its volatility, the effect of the latter being about one third as large as of the former.[4]

We focus on Italy during the years 1881 to 1912 for three reasons. First, Italian pre-WW1 migration represents one of the most significant voluntary, mainly economic-driven, population movements in history: more than 11 million Italian migrants left Italy from 1870 to WWI (ISTAT, 2012), with only Russian Jews being more likely to migrate.[5] Second, Italian international out-migration presents an extraordinary level of – largely unexplained – spatial and temporal heterogeneity across provinces, in terms of destinations and magnitudes (Fauri, 2015). Finally, the consequences of agricultural commodity market integration on migration are more likely to materialize in developing economies largely dependent on the primary sector. Indeed, about two thirds of the Italian labor force was employed in agriculture at the beginning of the 20th century (Felice, 2018). Moreover, Italian real wages were significantly lower than the

[2] Throughout the paper we use the term globalization – or first globalization era – as shorthand for the general process of labor movements and commodity market integration across the Atlantic Ocean.
[3] Symmetrically, the phenomenon reversed in the face of sharp price plunges. Harvey et al (2017) documented a positive linkage between commodity prices and short-run incomes.
[4] In theory, we might expect that volatility would have affected seasonal migration more than gross migration, but we cannot separate migrants into temporary and permanent categories in the analysis so we cannot test this. Hatton's (1995) model of migration explores the role of uncertainty (in employment) as an emigration determinant and data for UK outflows suggest that volatility explains short-run fluctuations more than long-run trends.
[5] Spitzer (2016) documents that, as of 1897, one third of the entire Russian Jewish population had emigrated.

European average and roughly equivalent to a fourth of those in the US (Williamson, 1995). Anecdotally, a series of government inquiries (Sonnino and Franchetti, 1877; Jacini, 1884; Faina, 1907) offered a dismal depiction of agricultural laborers across the peninsula. Often illiterate and exposed to malarial illnesses, most farmers lived in near-subsistence conditions.

The period 1881-1912 is of particular interest because it is part of the broader age of convergence between 1850 and 1914 which O'Rourke and Williamson's *Globalization and History* attributes to the "open economy forces of trade and mass migration" (O'Rourke and Williamson, 1999, 5). The era was also characterized by significant commodity price swings, somewhat masked by an overall downward trend for most commodities, as Asian silk, Indian rice and Russian and American cereals flooded the European market for the first time (O'Rourke, 1997). Such swings created what Diaz Alejandro (1982) has called the "commodity lottery", whereby price movements generated winners and losers as globalization forces increased exposure to shocks in the international market and often made terms of trade less favorable for peripheral and developing countries.

Our paper contributes to the literature on migration at large, with a focus on agricultural commodity market integration as a key determinant. We construct emigration rates from Italian provinces and examine the role of agricultural-price shocks as a driver of migration. Given the substantial variation in migration across Italian provinces, the focus on national data may have obscured the factors driving differential rates.[6]

A handful of papers focus on the role of income shocks in historical migration decisions. Spitzer (2016) analyses the impact of temporary income shocks – proxied by changes in agricultural yields – on Jewish migration from Russia. He shows that economic shocks might explain the timing of migration but underlying forces such as wage differentials and networks remained the key driving factors. Abramitzky et al. (2012, 2013) focus on Norwegian migration to the US and exploit variation in expected inheritance based on birth order to show that wealth was negatively correlated with migration, arguing that self-selection was negative from urban areas. Persaud (2017) focused on income volatility (through volatility in rice prices) as a push factor in Indian indentured servant migration.

[6] Moretti (1999) and Ardeni and Gentili (2014) point out the significant provincial variation that is usually ignored. Spitzer and Zimran (2018) also present anthropometric evidence that, while negative at the national level, the degree of self-selection of Italians coming through Ellis Island varied across provinces, being more positive in the South.

Turning to primary sources, we find further evidence that potential Italian migrants were liquidity constrained during our period. Coan (1997) presents interviews with some Italian-Americans who entered through Ellis Island. Mario Vina did not come to the US until two years after his father because the remittances received from the father were used for a sick relative. He describes how, "*In 1909, the year we came, my father got smart. He sent no cash, just the paid tickets*", (Coan, 1997, 38). Others describe vividly the role of networks and prepaid tickets in facilitating migration for poor Italians — Peter Mossini left Sicily in 1921, despite his father coming for work to the US from 1901-1912. The money that he sent home sometimes covered a ticket for one of his children, but Peter had to wait for a sibling to succeed sufficiently in the US to pay for his passage, and he stayed with family when he eventually migrated (Coan, 1997, 45).

## 2. DATA

Our dataset is the result of a significant digitization effort.[7] We combine annual information about international emigration and population at the province level with agricultural production structures along with agricultural commodity prices and landholding systems.

*EMIGRATION DATA*

Provincial emigration data are only available in absolute terms and reconstructing annual migration rates required the digitization of population series – to be found in yearly publications made available by ISTAT (*Annuario Statistico Italiano, ad annum*, since 1881).[8] Figures are taken from the *Annuario statistico dell'emigrazione italiana dal 1876 al 1925* (1926), a comprehensive volume that contains the universe of emigration statistics produced by the two main governmental migration agencies operating during our reference period: The *Direzione Generale di Statistica* (henceforth, DGS) and the *Commissariato Generale dell'Emigrazione* (henceforth, CGE), active since 1901.[9] We improve on the existing literature by employing a longer time series, 1881-1912, at a significantly higher level of disaggregation, *i.e.* the 69 Italian

[7] The data have been deposited here: http://doi.org/10.3886/E110063V1.

[8] Inter-censal population data are available on a yearly basis only after 1896. We conduct some robustness checks using population figures taken from the 1881 Census only. The results are presented in Table A12.

[9] The two bodies operated autonomously and developed different, often contradictory, definitions of migrants. They also used different data collection techniques. Such inconsistencies help explain why, despite similar overall trends, there might be discrepancies in absolute numbers. See Tortorici (2017) and Bevilacqua (2001) for an in-depth treatment of the matter.

provinces rather than the 16 regions. The most closely related studies are Gomellini and Ó Gráda (2013) – who carry out a regional analysis over a similar time frame – and Hatton and Williamson (1998) – who employed provincial data for only two census years, 1901 and 1911.

We focus on the DGS series as they are available since 1876 – unlike the CGE figures, which are only available from 1902 onwards. According to Italian law, would-be migrants had to undergo a two-step procedure to leave the country. First, they had to apply for a *nulla osta*, a document certifying their eligibility for a passport, which was then issued by local authorities. Second, they had to collect the actual passport, which was valid for 3 years. DGS series are based on the number of *nulla osta* issued by each municipality, aggregated to the province level.

These data exhibit two caveats. First, would-be migrants could decide not to leave and, in rare cases, local authorities could reject passport applications. Conversely, several destination countries allowed European migrants to enter without passports, decreasing incentives to follow standard legal procedures in sending countries (Keeling, 2013). The DGS series thus measures emigration with error, although there is no reason to expect this to vary systematically across provinces. Second, the DGS series only differentiates between European and overseas emigration at large, masking destination-specific patterns as well as any gender, age, and occupational dimensions. The occupational decomposition of flows is only available at the national level and confirms that about 70% of Italian passport applicants were employed in the agricultural sector. In light of this piece of evidence, our analysis is likely to capture the main driving forces of provincial out-migration.

Our paper shares a limitation with the existing literature as it is not possible to account either for return nor for seasonal migration. Indeed, recent contributions (Bandiera et al., 2013) have highlighted how considerable these phenomena were during the era of mass migration. Since we are interested in investigating how a changing economic environment might influence migration at large, such constraints should not invalidate our analysis despite inevitably introducing some error in the measure of migration stocks at destination.

We complement the provincial emigration data with Ellis Island administrative records, a comprehensive collection of individual-level data about the universe of Italian passengers setting foot on Ellis Island, the main immigration hub in the United States. We focus on Italian passengers between 1892 and 1912 – a grand total of about 3.5 million data points. Although these data have been used in recent studies,[10] we innovatively employ them to study the determinants of Italian emigration to the US at the level of the province of origin. In particular, we use the last municipality of residence of each individual to identify their province of origin. This required extensive data cleaning as missing or misspelled localities were common. We designed an ad-hoc script that groups municipalities by spelling mistake and confronts each string with an all-encompassing list of Italian municipalities – including those that changed denomination, were aggregated or split. After identifying both perfect and imperfect matches, the script returns a set of suggestions based on string distance. Finally, we manually select the most appropriate match (Tortorici, 2017). This process allows us to retrieve about 1.62 million observations – about 53 per cent of Italian migrants who reported their last residence.[11] This largely explains the differences in sample sizes across the columns in our results Tables 2-5 below, as the Ellis Island series began only in 1892 and represents a lower bound estimate of out-migration because we could not match all migrants to their place of last residence.

*[Figure 1 here]*

The emigration data presented in Figure 1 show a significant degree of variability across provinces, regardless of destination choice. Emigration rates approach zero for some Southern provinces in the early 1880s, while maxima are driven by the provinces of Udine and Belluno, in the North-East. These areas experienced intensive emigration towards European countries – even compared with the later emigration rates of Southern provinces towards the US.[12] As mentioned above, emigration rates obtained from Ellis Island administrative records can be considered as a lower bound due to attrition in the matching procedure. Table 1 presents summary statistics for the variables used in the analysis. Migrant networks are defined as the one-year lagged sum of absolute migration from the DGS series.

*[Table 1 here]*

[10] See, for example, Bandiera et al. (2013), Spitzer and Zimran (2018), and Ward (2017).
[11] For completeness, note that information on last residence is particularly coarse before 1901. Our analysis might therefore underestimate emigration to Ellis Island throughout the 1890s.
[12] Before 1900, Europe and South America were popular destinations, while after 1900 the US become increasingly dominant – despite a sizable stream of Italian migrants still opting for Europe.

*GLOBAL PRICE EXPOSURE INDEX AND OTHER VARIABLES*

In order to investigate the impact of agricultural commodity price shocks on international out-migration from Italian provinces, we construct a global price exposure (GPE) index that captures provincial exposure to global price movements across relevant agricultural commodities produced in Italy during the sample period. We construct this measure annually for the period 1881 to 1912.

We proxy international commodity prices with cif import prices, digitized from the *Annuario Statistico Italiano*.[13] This choice is motivated by endogeneity concerns, to ensure that commodity prices are exogenous to the economic activity of each province. Italy was a small, poor economy during the nineteenth century, and thus was not a price maker in international commodity markets. Furthermore, we note that much of the effects of changes in trade policies would be absorbed by the year fixed effects we employ in all specifications, as they are enforced nationally.

Figure A1 shows that the productive structure of Italian provinces was quite diverse – *i.e.* different crops were produced in different areas – and fluctuations in, say, wheat prices might not affect provinces equally. In order to account for such heterogeneity, we weight price movements using purposely-digitized information about crop allocations in terms of acreage, averaged over the period 1876-1881.[14] Our measure incorporates all major agricultural commodities produced in Italy: wheat, corn, oats, wine, olive oil and rice. Data have been digitized from the *Annuario Statistico Italiano* (1886, ASI henceforth).[15] We use acreage rather than output because many authors have expressed concerns about the reliability of historical Italian output data.[16] However, in the interest of completeness, we also construct a revised GPE index, which uses output rather than acreage shares. The results are consistent with the main results in this paper and are presented in Table A1 of the Online Appendix.[17] Following the recent literature, we keep the

[13] The price series covers the period from 1881 to 1912. We test for the nonstationarity of the price series using the Augmented Dickey-Fuller test.
[14] As an alternative specification, we construct a GPE index based on crop shares averaged over the period 1879-1883. Table A2 in the Online Appendix presents the results using this alternate GPE definition. We also construct a GPE measure based on crop shares, averaged over the period 1891-1894 (Table A3 in the Online Appendix).
[15] To the best of our knowledge, there is no systematic recording of acreage data at the province level after 1894. Data are available as 1876-1881 averages (ASI, 1887-1888) or yearly measures for 1891 and 1894 (ASI, 1892, 1895).
[16] See, for example, Fenoaltea (2011) who argues that the agricultural crisis of the 1880s might actually be attributed to misreporting.
[17] The output data are averages for 1876-1881 and were taken from the *Annuario Statistico Italiano* (1886).

share of land assigned to each crop fixed to the beginning of our time frame in order to avoid potential endogenous re-allocation issues.[18] Note also that provincial crop mixes did not shift significantly over time (see Figure A2 in the Online Appendix).

Our Global Price Exposure (GPE) index is constructed as follows:

$$GPE_{it} = \sum_{j=1}^{J} \gamma_{ji,76-81} P_{jt} \quad (1)$$

where $\gamma_{ji,76-81}$ represents the 1876-1881 average share of crop *j* farmed in province *i*, and $P_{jt}$ is the import price of crop *j* at time *t*. The use of the pre-period crop shares, $\gamma_{ji,76-81}$, mitigates concerns about the potentially endogenous response of local crop production following price changes. The GPE index reflects provincial exposure to agricultural-price changes over time and is in line with the recent literature that examines the effect of competition from China on labor markets following its accession to the WTO in 2001.[19] It is similar to the measures of local labor market exposure to import competition developed by Topalova (2007) and Autor et al. (2013). It is further in the same spirit as the composite commodity price indices that have been constructed as weighted averages of international prices and analyzed in the macroeconomics literature (see, for example, Harvey et al (2017) which expanded and improved on the well-established Grilli-Yang series).[20] We argue that international prices, which our measure proxies for, are driven by international demand and supply forces and are independent of changes in these factors at the province level. We expect that changes in supply costs dominated during these years, given the advances in shipping technology.

Figure 2 displays a clear downward trajectory of international real prices (in 1912 *lire*) of the six major agricultural commodities farmed in Italy up to the mid-1890s. Prices of the dominant crop, wheat, reached their *nadir* in 1894, at about half their 1879 level. Prices partially recovered thereafter but exhibited evident year-to-year fluctuations. Italian wheat exports dropped to zero while imports followed an upward trend.

[18] We use average shares as yearly measures might be influenced by crop rotations. We acknowledge that intercropping might have been a widespread practice. However, due to data availability, we are only able to focus on the main crops. Recent contributions to the migration literature often account for only one crop, as in Bazzi (2017) and Persaud (2017). As mentioned above, we experimented with using output shares to compute the GPE measure, to deal with the problem of intercropping and additionally used alternative acreage shares, as shown in the Appendix Tables A2 and A3.

[19] See also, among others, Pierce and Schott (2016) and Acemoglu et al. (2016).

[20] These series typically look at a wider range of commodities, adding oil and metals to agricultural staples.

*[Figure 2 here]*

We argue that these price movements were perceived by Italian farmers as a result of agricultural commodity market integration, although food prices are likely to have declined slightly.[21] The measure is admittedly broad and might capture variations in global supply and demand that could have driven agricultural-price changes: we therefore interpret our GPE index as a proxy for general commodity market trends.

In addition to the GPE measure, we also construct a measure of price volatility, in line with Persaud (2017). Analyzing the determinants of Indian indentured servitude, he argues that volatility in the price of rice, the dominant crop, could be thought as a proxy for income uncertainty, a potential push-factor driving migration.

More formally, both unfavorable and uncertain prices for agricultural goods could reduce the demand for farm labor by lowering the expected marginal value product of workers, thus reducing their wage. In addition, low and variable prices will reduce marginal profits for small land-holders who may be forced to sell or abandon their plots, or encourage temporary migration to finance buying more land to compensate for falling prices (Cinel, 1991, 166). Both of these channels will increase the incentives for affected workers and land-holders to migrate.[22]

For each crop *j* and time period *t*, price *p* volatility is computed over 5-year windows; it corresponds to the standard deviation of logged inter-temporal import price ratios. The GPE volatility index is then constructed similarly to the GPE presented in equation (1), using average acreage shares at provincial level interacted with this measure of volatility in agricultural commodity prices for each crop:[23]

$$Volatility_{i,t} = sd\left[\ln\left({p_{j,t-h+1}}/{p_{j,t-h}}\right)\right] \; with\; h = 1, \ldots, 5 \qquad (2)$$

[21] Indeed, there is evidence of rising living standards and consumption from the 1880s onwards (Fenoaltea, 2011). This is not in contrast with our hypothesis because it relies on the idea that out-migration increased following the weakening of liquidity constraints, not because of increased hardship.

[22] Barkan (2013) documents how small-holders overcame liquidity constraints by selling land to finance migration.

[23] We also compute volatility by varying the time-window considered, from 3 to 7-year windows. The results using these alternative volatility measures are consistent with our findings and are presented in the Online Appendix (Tables A5, A6, A7, A8). The trade-off in deciding on the appropriate time-window is between sample size reduction and having a sensible measure (volatility over 2 years would not be very informative).

Our extended specifications control for factors that have been found important in other studies, such as Hatton and Williamson (1998) — migrant stocks and landholding structures. We include migrant stock at destinations as they might significantly lower the psychological, and financial costs involved with transnational migrations through different channels, thus alleviating liquidity constraints to some extent. These include information about job vacancies and local information about customs and lodgings within immigrant communities.

We further collected information about landownership from the 1881 Italian Census. The share of non-landowners employed in the agricultural sector varies across provinces, with lower values to be found in North-Western Italy. Overall, the mean is quite high, with about 85 per cent of those employed in the primary sector not owning land (Table 1).

Finally, we include a measure of industrialization, as better employment opportunities in the secondary sector might represent an alternative to migration for would-be migrants employed in agriculture. We measure provincial industrialization using the recent estimates by Ciccarelli and Fenoaltea (2013). Industrialization varied significantly across Italian provinces with the *maximum* being Milan, and the minimum Sassari – a Northern province of Sardinia. Overall, aside from Naples, the South was on average less industrialized than the North.

## 3. ECONOMETRIC SPECIFICATION

Our estimation relies on the idea that international price fluctuations transfer to local economies and affect international out-migration through consequential income shocks. For this to hold, two conditions must be satisfied. First, agricultural commodity prices should be positively related to agricultural incomes. Second, local and global agricultural markets should exhibit a significant degree of integration. Both assumptions are supported in the historical literature.

Italy remained a fairly open economy and progressively integrated into the global market throughout the entire time span of reference. However, rapidly increasing competition on traditionally land-intensive crops, *e.g.* wheat and corn, alarmed landowners who envisaged reduced domestic market shares. Even though protectionist measures remained mild until the early-1880s, Italy levied a series of successive tariffs on wheat from the mid-1880s – with the tariff corresponding to 40 per cent, *ad valorem*, by 1913. In general, overall agricultural protection remained relatively low, with the exception of a short spell in the 1890s. Italian agricultural trade policies did not define it as an unusually protectionist economy in the European context. Indeed, there is evidence that tariffs did not generate major compositional shift in agricultural mixes, while possibly increasing wheat production (Federico and O'Rourke, 2000). Overall, protectionism did not prevent prices from falling across Europe as technological change and market integration acted in the opposite direction. This point is supported by evidence of wheat prices co-moving within Italian major markets and across Europe.[24] As explained in the previous section, our GPE index is based on cif import prices, weighted using crop shares in terms of acreage.

The main econometric specification is as follows:

$$\log(migration\_rate_{it}) = \beta \log(GPE_{it}) + X'_{it}\gamma + \delta_i + \vartheta_t + \varepsilon_{it} \qquad (3)$$

where $migration\_rate_{it}$ is the out-migration rate from province $i$ at time $t$ and $\beta$, the key coefficient of interest, captures the effect of agricultural commodity market integration through price shocks. Note that both the migration rates and the GPE index are logged, so $\beta$ can be interpreted as an elasticity. We run three separate sets of regressions, taking into account different macro-area destinations: Europe; Mediterranean and overseas destinations at large (Canada, US, Latin America and Oceania) and the United States.[25] Equation (3) includes time-varying characteristics at the province level ($X'_{it}$), namely industrialization, land ownership and migration networks. The specification also includes province and time fixed effects (respectively, $\delta_i$ and $\vartheta_t$). Standard errors are clustered at the province level.

Further, we interact *GPE* with the provincial share of non-landowners[26] in 1881 because globalization might have had differential effects depending on what landholding structure prevailed. Additionally, we estimate specifications that include an interaction between our *GPE* and a dummy variable for Southern provinces to account for North/South economic differences.

Finally, in the most stringent specification, we include region by time dummies: this set of fixed effects controls for any change at regional level over time which may affect our results, such as differences in financial development or any other type of time-varying characteristics

[24] See Figure A4 in the Online Appendix for more details.
[25] We conducted a similar analysis using internal rather than international migration rates, which are available after 1906. We found no statistically significant effect of the GPE index on movements across Italian provinces. Overall, international migration flows dwarfed internal migration. Results are reported in Table A13 in the Online Appendix.
[26] This class includes sharecroppers, day laborers, salaried laborers as well as renters.

across Italian regions.

## 4. RESULTS

How did globalization – thought of as agricultural commodity market integration – influence international emigration choices from Italian provinces between 1881 and 1912? Table 2 presents baseline estimates of the impact of the *GPE* index on provincial emigration rates towards different international macro-areas. Columns 1-4 pertain to, respectively, total overseas; transatlantic; US (Ellis Island); and European out-migration rates.

*[Table 2 here]*

We find that the GPE index has a statistically significant and positive effect on emigration rates, with transoceanic migration being the most responsive.[27] The log-log regression results show that the effect of a 1% increase in GPE ranges from a 0.64% increase in the Ellis Island out-migration to a 1.72% increase in transoceanic migration. Given that our agricultural-price index is essentially a weighted average of import prices – with weights reflecting initial provincial crop intensities – we interpret this coefficient as indicating that emigration increased in periods when international prices increased. The rationale behind this mechanism relies on the idea that agricultural incomes – especially those of small land-holders – increase proportionally to agricultural commodity prices. Conversely, exogenous price plunges are expected to have a negative effect on agricultural incomes and prevent would-be migrants from migrating as liquidity constraints bind. Our finding is in line with Bazzi (2017), who also explored the relationship between income and propensity to migrate in contemporary Indonesia and finds an inverted U-shaped relation between wealth and migration. Throughout our sample period, Italy remained a slowly-developing country where liquidity constraints prevailed despite negative price shocks making migration more appealing. This result is supported by Faini and Venturini (1994) who argue that Italy was indeed caught in a poverty trap that curbed out-migration. Spitzer and Zimran's (2018) paper also suggested that migrants from poorer Italian provinces and those who were able to finance their own trips were more positively selected, consistent with

[27] Results are robust to lagging the GPE measure for one or two years, and the coefficient increases in size, takes indicating that it may take time to respond to the price shock. Results are available upon request.

the presence of significant liquidity constraints.[28] Evidence of liquidity constraints have further been demonstrated in data on selection of migrants in Covarrubias et al (2015), who showed that for the period 1899-1932 across 39 countries that sent migrants to the US, when GDP increased, the average skill level of migrants decreased. Our results also fit with theory from Ciccarelli et al (2018) who present a version of their model with liquidity constraints which shows that, in that case, areas with initially low wages could respond to increasing real incomes by migrating—they believe that this effect might be magnified in areas with more concentrated landholding or monopsonistic labor markets. Finally, Spain and Italy are often analyzed together as development and migration laggards. Sanchez-Alonso (1995, 2000) found that agricultural wages and migration rates were positively correlated at the province level for pre-1914 Spain.

Table 3 expands our baseline specification and includes a measure of *GPE* volatility over a 5-year window. We add this independent variable because would-be migrants might have been influenced by price – and therefore income – uncertainty, over and above their response to changing income trends. Indeed, we find evidence that this was the case. The coefficient is statistically significant both in columns 1 and 2 – for aggregate and transoceanic emigration rates – while there seems to be no influence on emigration towards the US (column 3) and to other European countries (column 4).[29] Overall, although it is safe to argue that uncertainty might have driven emigration – which could then be thought of as a coping strategy – our results do not offer any insights on how price volatility might affect individuals differently across the income distribution. For instance, Persaud (2017) presents evidence that rice price volatility – the major crop in India, where the study is set – was a key driver of indentured emigration especially for lower status individuals. A possible interpretation for the fact that our volatility coefficient is not significant across all specifications could be seasonal migration – which we cannot capture – towards both North and South America.

*[Table 3 here]*

[28] Wegge (1998), looking at the mid-nineteenth century, found that liquidity constraints prevented the poorest laborers from German villages from migrating, and led to the middle class dominating migrant ranks.

[29] We obtain similar results when including the GPE volatility measure by itself. Results are available upon request.

Table 4 presents our main specification. It includes a full set of controls, *i.e.* provincial industrialization, migrant stocks at destination and landholding structures.[30] Industrialization is not significant except for migration towards the US, where it has the anticipated negative sign, indicating that more industrially developed provinces were able to absorb the surplus of agricultural labor, thus decreasing the propensity to migrate. These coefficients are in line with the fact that migration shifted southwards at the turn of the 20th century. Indeed, Southern provinces were comparatively less industrialized than Northern ones and the secondary sector could not absorb agricultural labor surplus to any significant extent. As expected, migrant networks – despite their likely upward bias because it does not discount return and cyclical migration – have a crucial role in explaining provincial emigration rates, as existing migrant stocks at destination facilitated chain migration through financial, remittances, and information channels (see Moretti, 1999; McKenzie and Rapoport, 2010; and Gomellini and Ó Gráda, 2013). This result is consistently statistically significant at the 1 per cent level and similarly sized across all destinations we consider. In addition, Table A1 in the Online Appendix presents the results of estimating this specification using an alternative measure of GPE that was constructed using output rather than acreage shares. The results are largely consistent with the preferred specification discussed above.

*[Table 4 here]*

The analysis provided above shows the role of commodity price changes and uncertainty in affecting provincial migration patterns. It is, however, silent about potentially different effects across land-holding structures. We explore the role of land tenure – also heterogeneous in Italy – in Table 4, which includes an interaction term between our *GPE* index and the share of non-landowners at provincial level, as measured in 1881. The estimated coefficient is negative, indicating that small owners' incomes might have been more responsive to price fluctuations. On the other hand, landless laborers might have benefited from food price reductions, being able to count on a higher income. Another possible mechanism might be that higher agricultural prices

[30] Table A11 in the Online Appendix presents the estimation results including welfare ratios from Federico et al (forthcoming). Note that we do not control for wages in our main specification, for two reasons. First, the provincial series only began in 1905 and second, wages are likely to be endogenous to agricultural or industrial business cycles.

increased incentives to invest in agriculture: given that temporary and seasonal migration had remained quite high throughout this period, small plot-owners' decision to migrate may have been motivated by a desire to acquire extra capital.

According to Hatton and Williamson (1998), differentials in North/South migration patterns might partly be due to the relative abundance of non-agricultural labor opportunities in relatively more industrialized Northern areas, or the relative underemployment of agricultural labor and tenants in the South, augmented by the generally low level of development. Table 5 presents the results of a specification that includes an interaction term between our GPE index and a dummy variable for Southern provinces: if there is a non-linear relationship between income and migration, we might expect to see a differential effect in these areas. We find that there is a significant difference when comparing transoceanic emigration rates at large (column 2) to emigration rates to the US (column 3). The effect of globalization on transoceanic emigration rates is lower in Southern provinces, but we find the opposite result when we look at emigration rates towards Ellis Island. Indeed, the estimated effect of the GPE index doubles in Southern provinces in the latter case. This finding is consistent with the overall migration patterns discussed in Section 2: emigration centers gradually shifted from Northern to Southern provinces while preferences converged on the US. This result suggests that liquidity constraints on migration bound more tightly in Southern provinces, which were less developed and had smaller migrant stocks at destination throughout the 1880s and 1890s.

*[Table 5 here]*

In terms of magnitude, we focus on the coefficients on global and transoceanic migration. These are the most stable coefficients across our specifications, which is reasonable given that the Ellis Island migration is measured with error and can only be used from 1892 onwards, and European migration is most likely to be compromised by our inability to account for return migration. The coefficients on global emigration range from 1.33%-3.66% across Tables 2-5, suggesting an elastic response of migration to all destinations as a result of agricultural commodity price increases. If we take the sub-period from 1893 to 1900, we see that the GPE index increased by almost a standard deviation, by 18.26%. Taking our 1.33-3.66 coefficient range seriously, this implies a 24.29-66.83% increase in the global migration rate. The actual increase in global migration across provinces over that 8-year span was 138.6%, so it appears that commodity price trends can explain almost 50% of the substantial increase in average migration.

## 5. ROBUSTNESS CHECKS

We conduct a series of additional robustness checks, which we present in the Online Appendix. As mentioned in Section 2, we calculate the GPE index on the basis of provincial acreage crop shares averaged over the period 1876-1881. Tables A2 and A3 present the results of the full specification when we construct the GPE index on the basis of 1879-1883 and 1891-1894 acreage respectively. Results are not sensitive to different ways of computing acreage shares.

In the analysis presented so far, we use import price series rather than domestic agricultural commodity prices to preserve the exogeneity assumption. Indeed, one may be concerned that there may be factors jointly impacting both domestic prices and out-migration rates, such as a bad harvest or harsh weather conditions. As a robustness check, we replicate the analysis using the national agricultural commodity price series from the *Sommario di Statistiche Storiche 1861-2010*.[31] The results, presented in Table A4, support our main findings. This is indeed not surprising: Figure A3 in the Online Appendix presents the international and national wheat price series. The two series display a very high degree of correlation, suggesting that markets were indeed integrated and price trends mainly driven by global technology shocks. Furthermore, Figure A4 shows that local (major-city-wise) wheat price series – purposely digitized from the *Movimento dei prezzi di alcuni generi alimentari dal 1862 al 1885* (Direzione Generale della Statistica, 1886) – co-move at least since the early 1860s, suggesting that there was considerable integration across internal markets.

Next, we measure volatility over different time windows, from 3 to 7 years. We present the results of this exercise in Tables A5, A6, A7 and A8. Results are largely unchanged.

We also split the sample into two periods: pre-1894 and post-1894 in order to capture the effect of the GPE index when agricultural commodity prices were generally trending downwards (pre-1894) or upwards (post-1894). Results are presented in Tables A9 and A10. The estimated coefficient of the GPE index is positive and statistically significant in both cases and its magnitude is larger in the former period, suggesting that incentives did play an important role

[31] We use production prices for all the crops apart from the wine series, for which we use consumption prices.

and liquidity constraints were stronger earlier on. A price increase when prices are generally spiraling downwards matters more than a price increase when prices have already recovered.

We run an additional robustness check including welfare ratios as an additional control variable, generously provided by Federico et al. (forthcoming), which are available since 1905 only. Results are presented in Table A11: we do not find any statistically significant effect of welfare ratios on out-migration rates, possibly because the series is relatively short and migration networks were already fully functioning, thus decreasing the sensitivity of migration to wage differentials – which were in any case falling gradually.[32]

We further try an alternative way of computing migration rates. As already mentioned, pre-1896 population data are not available for all years which means that some observations are dropped in the early part of the sample. As a robustness check we construct out-migration rates using provincial population figures taken from the 1881 Census only. Table A12 presents the results of this alternative measure. They are in line with the previous findings.

Finally, Table A15 presents the results of a more demanding specification where we include region by time dummies, which would control for any change at regional level over time, to assuage any remaining concerns about changing endogenous tariff policies or other time-varying changes, such as differences in financial development across Italian regions. The results are in line with the previous findings.

## 6. CONCLUSION

Analyses of Italian international migration flows at the sub-national level have been relatively limited in the historical literature on the Age of Mass Migration. We fill this gap by exploiting the extraordinary heterogeneity across Italian provinces in terms of out-migration rates and agricultural production structures.

This paper revolves around exogenous shocks to agricultural commodity prices – and therefore agricultural incomes – that occurred in Italy as a result of the first wave of

[32] As a further robustness check, we introduce a specification that includes a wage ratio which is available for the United States only, hence restricting the focus to the emigration rate to the US (Ellis Island). The provincial welfare ratios from Federico et al. (forthcoming) are only available for 27 out of 69 provinces. We divide these figures by the commonly used real wage series described in Williamson (1995), Appendix 1. Overall, the estimated coefficients presented in Table A14 are in line with the main ones presented in Table 4. The coefficient on GPE is still positive, but it is less precisely estimated due to the smaller sample size, which is also not representative of Italy as a whole.

globalization. We use start-of-period crop mixes to assign differential treatment intensity to each province and explore the effect of price fluctuations and uncertainty on migration flows. We find that higher prices translated into higher agricultural incomes and therefore higher migration rates; this mechanism suggests that would-be migrants faced binding liquidity constraints until their income reached a given threshold.

This article represents an opportunity to compare the current globalization process with its first historical counterpart, drawing lessons for today's developing countries by analyzing then-developing Italy. Indeed, as transport and communication technology continues to advance and more countries join international trade agreements, the consequences of market integration will likely become increasingly salient. Today, many developing countries are vulnerable to commodity price shocks and exhibit even greater wage differentials with potential destinations. However, stringent immigration laws prevent most would-be migrants from embarking on international migration. Our study provides suggestive evidence of what might occur should developing countries gain more access to migration during their ongoing process of globalization.

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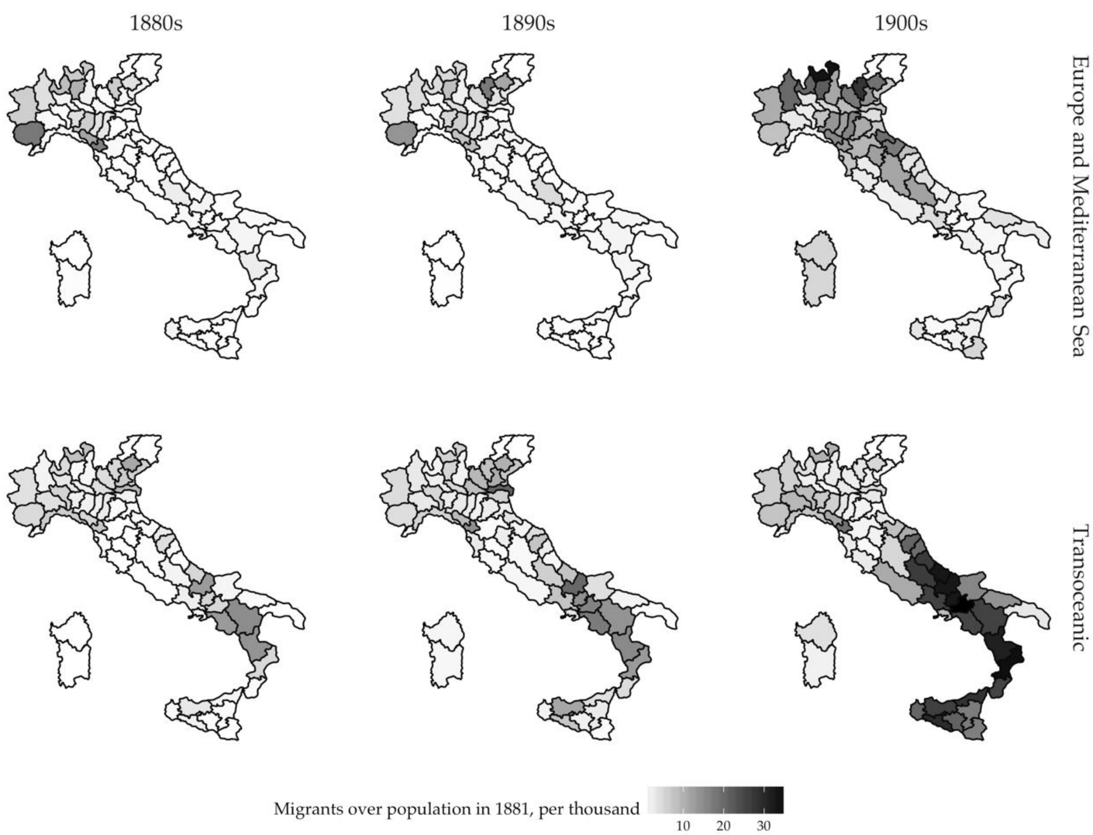
**FIGURES AND TABLES**

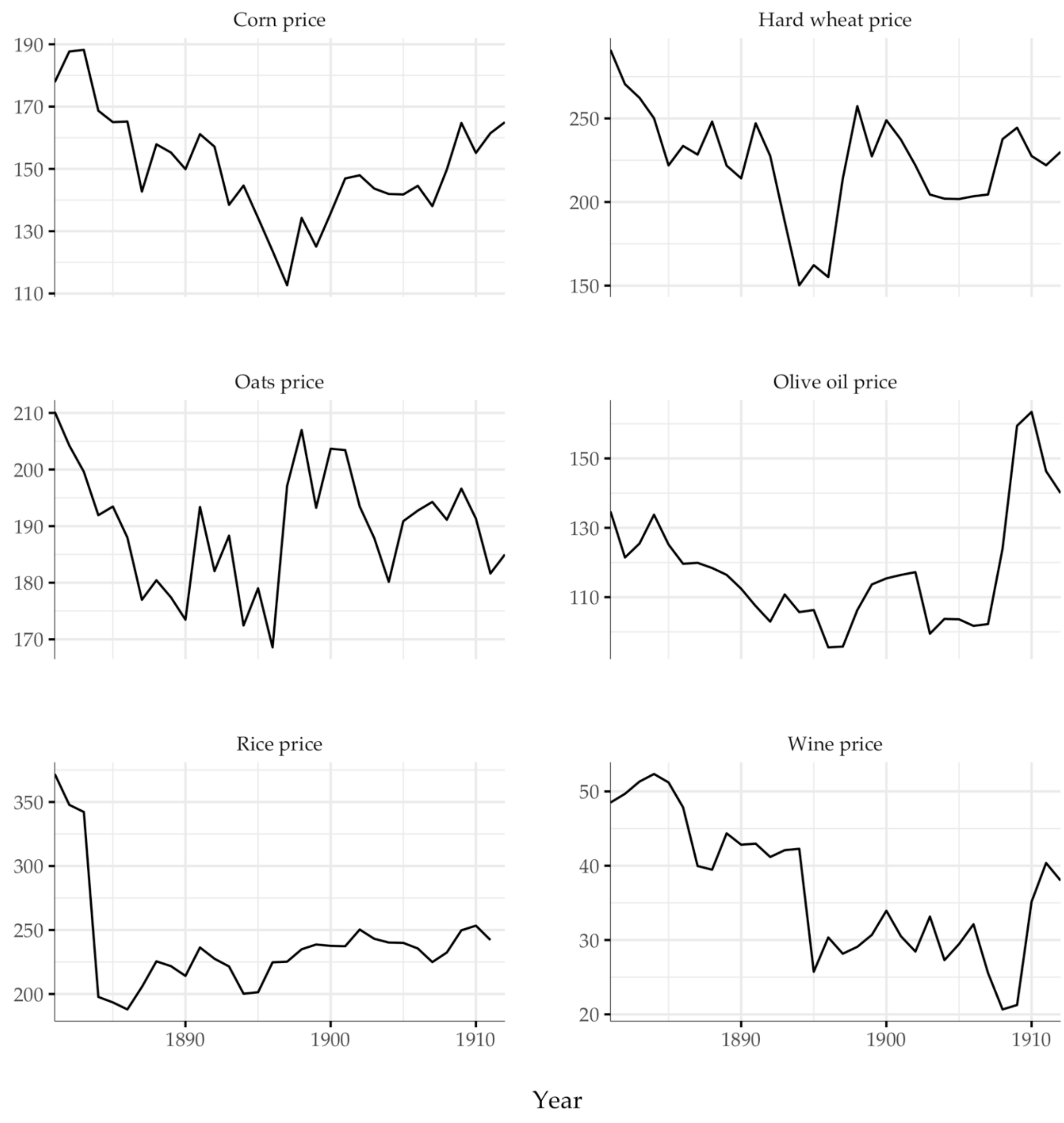
**FIGURE 1: PROVINCIAL EMIGRATION RATES[33]**




*Source: Annuario Statistico dell'emigrazione italiana dal 1976 al 1925.*

[33] Note that these maps do not include the provinces of Udine and Belluno because emigration rates from both areas were high enough to mask the actual cross-province variation elsewhere. Udine and Belluno provinces are included in the analysis. Estimation results are also robust to their exclusion.

**FIGURE 2: INTERNATIONAL CROP PRICES**




*Source: Annuario Statistico Italiano*

**TABLE 1: SUMMARY STATISTICS**

| Variables | Mean | SD | Minimum | Maximum |
|---|---|---|---|---|
| GPE (logged) | 2.26 | 0.02 | 2.18 | 2.31 |
| GPE, 5-year volatility (logged) | 1.91 | 0.06 | 1.65 | 2.04 |
| European emigration rate per 1000 | 6.95 | 13.63 | 0 | 137.38 |
| Global emigration rate per 1000 | 14.85 | 16.32 | 0 | 143.75 |
| Transoceanic emigration rate per 1000 | 7.99 | 10.02 | 0 | 58.68 |
| European migration network | 29,094.44 | 87759.11 | 0 | 1106798 |
| Global migration network | 62,131.3 | 106544.89 | 0 | 1228871 |
| Transoceanic migration network | 33,373.53 | 50096.87 | 0 | 324715 |
| Ellis Island (US) migration network | 4,768.44 | 13511.42 | 0 | 155449 |
| Ellis Island (US) emigration rate per 1000 | 2.74 | 4.27 | 0 | 26.12 |
| Population (1881 census) | 412,458.67 | 222,663.86 | 114,295 | 1,114,991 |
| Yearly population | 466,499.81 | 269,462.58 | 114,295 | 1,743,723 |
| Share of non-landowners | 84.43 | 10.25 | 41.11 | 97.32 |
| Provincial industrialization (Ciccarelli and Fenoaltea, 2013) | 0.91 | 0.34 | 0.43 | 2.26 |

**TABLE 2: GLOBALIZATION AND PROVINCIAL EMIGRATION RATES**

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| Variables | *Global emigration rate* | *Transoceanic emigration rate* | *Ellis Island (US) emigration rate* | *European emigration rate* |
| GPE | 1.3324* | 1.7217** | 0.6433*** | 0.7187 |
| | [0.668] | [0.854] | [0.238] | [0.613] |
| Observations | 1,442 | 1,432 | 1,115 | 1,423 |
| No. of provinces | 69 | 69 | 69 | 69 |
| Adj. R-squared | 0.543 | 0.435 | 0.659 | 0.535 |

Notes: This table reports the estimates of equation (3) for all possible destinations. It includes our GPE measure without any additional controls. All regressions include province and year fixed effects. All variables are logged. Standard errors are clustered at the provincial level.
Significance: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

**TABLE 3: GLOBALIZATION, 5-YEAR VOLATILITY AND EMIGRATION RATES**

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| Variables | *Global emigration rate* | *Transoceanic emigration rate* | *Ellis Island (US) emigration rate* | *European emigration rate* |
| GPE | 1.6450** | 1.9399** | 0.6889*** | 0.8597 |
| | [0.732] | [0.872] | [0.222] | [0.708] |
| GPE, 5-year | 0.6570** | 0.6279* | -0.1508 | -0.1353 |
| volatility | [0.258] | [0.326] | [0.235] | [0.318] |
| Observations | 1,375 | 1,369 | 1,115 | 1,356 |
| No. of provinces | 69 | 69 | 69 | 69 |
| Adj. R-squared | 0.538 | 0.374 | 0.660 | 0.544 |

Notes: This table reports the estimates of equation (3) for all possible destinations. It includes both our GPE index and GPE volatility measures – based, respectively, on equations (1) and (2). All regressions include province and year fixed effects. All variables are logged. Standard errors are clustered at the provincial level.
Significance: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

**TABLE 4: FULL SPECIFICATION**

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| Variables | *Global emigration rate* | *Transoceanic emigration rate* | *Ellis Island (US) emigration rate* | *European emigration rate* |
| GPE | 3.6225** | 2.2614* | 1.1432 | 5.8793*** |
| | [1.409] | [1.222] | [1.048] | [1.667] |
| GPE, 5-year volatility | 0.4646*** | 0.4061** | -0.0707 | -0.0129 |
| | [0.171] | [0.183] | [0.166] | [0.246] |
| Provincial industrialization | -0.0208 | -0.5223 | -1.2911*** | 0.2606 |
| | [0.456] | [0.531] | [0.228] | [0.505] |
| GPE x share of non-landowners | -0.0404** | -0.0237 | -0.0085 | -0.0665*** |
| | [0.016] | [0.015] | [0.012] | [0.019] |
| Global migration network | 0.9342*** | | | |
| | [0.057] | | | |
| Transoceanic migration network | | 0.8653*** | | |
| | | [0.044] | | |
| Ellis Island (US) migration network | | | 0.2377*** | |
| | | | [0.048] | |
| European migration network | | | | 0.9898*** |
| | | | | [0.057] |
| Observations | 1,374 | 1,368 | 1,051 | 1,355 |
| No. of provinces | 69 | 69 | 69 | 69 |
| Adj. R-squared | 0.761 | 0.639 | 0.721 | 0.751 |

Notes: This table reports the estimates of equation (3) for all possible destinations. It includes both our GPE index and GPE volatility measures – based, respectively, on equations (1) and (2) – and an additional set of controls. Migrant networks do not account for return/temporary migration. Non-landowners include sharecroppers, day laborers, salaried laborers as well as renters. All regressions include province and year fixed effects. All variables are logged. Standard errors are clustered at the provincial level.
Significance: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

**TABLE 5: NORTH/SOUTH DYNAMICS**

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| Variables | *Global emigration rate* | *Transoceanic emigration rate* | *Ellis Island (US) Emigration rate* | *European emigration rate* |
| GPE | 3.6578*** | 2.2740** | 1.5377** | 5.8641*** |
| | [1.274] | [1.133] | [0.656] | [1.690] |
| GPE, volatility 5 years | 0.4300** | 0.3788** | -0.0660 | 0.0058 |
| | [0.172] | [0.188] | [0.159] | [0.247] |
| Provincial industrialization | -0.0635 | -0.5502 | -0.9885*** | 0.2837 |
| | [0.464] | [0.534] | [0.204] | [0.507] |
| GPE x share of non-landowners | -0.0368** | -0.0204 | -0.0207** | -0.0683*** |
| | [0.015] | [0.015] | [0.008] | [0.019] |
| GPE x South | -0.7971* | -0.6805* | 1.5745*** | 0.4164 |
| | [0.416] | [0.391] | [0.235] | [0.477] |
| Global migration network | 0.9348*** | | | |
| | [0.056] | | | |
| Transoceanic migration network | | 0.8691*** | | |
| | | [0.043] | | |
| Ellis Island (US) migration network | | | 0.2385*** | |
| | | | [0.044] | |
| European migration network | | | | 0.9911*** |
| | | | | [0.057] |
| Observations | 1,374 | 1,368 | 1,051 | 1,355 |
| Number of provinces | 69 | 69 | 69 | 69 |
| Adjusted R-squared | 0.762 | 0.640 | 0.739 | 0.751 |

Notes: This table reports the estimates of equation (3) for all possible destinations. It includes both our GPE index and GPE volatility measures – based, respectively, on equations (a) and (2) – and an additional set of controls. Migrant networks do not account for return/temporary migration. Non-landowners include sharecroppers, day laborers, salaried laborers as well as renters. The South dummy excludes Central Italian provinces. All regressions include province and year fixed effects. All variables are logged. Standard errors are clustered at the provincial level.
Significance: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$