Journal of Demographic Economics (2020), **86**, 435–477
doi:10.1017/dem.2020.7

RESEARCH PAPER

Migration-induced transfers of norms: the case of female political empowerment

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(Received 19 July 2018; revised 25 March 2020; accepted 27 March 2020)

Abstract
This paper empirically investigates the effect of transnational migrants on gender equality in the country of origin measured by the share of women enrolled in the lower chamber of National Parliaments. We test for a "migration-induced transfer of norm" using panel data from 1970 to 2010 in 10-year intervals. Total international migration has a positive and significant effect on female political empowerment in countries of origin conditional on the initial female parliamentary participation in both origin and destination countries. Endogeneity issues are taken into account and results are tested under specific geo-political subsamples.

Key words: Endogeneity; gender discrimination; international migration; panel data

JEL Classification: F22; J16; D72; C33

1. Introduction

In countries where women have minimal control over resources and a limited voice in decision-making, the exposure to new ideas through international migration can set off, facilitate, or catalyze gender parity [Hugo (2000)]. By looking at the shares of women enrolled in the lower chamber of National Parliaments, this paper explores the role of international migration as a measure of exposure to foreign practices related to female political empowerment.[1] During their stay abroad, migrants familiarize with different values, norms, and other forms of behaviors which are specific to the host country. They form networks of relationships and acquire knowledge on new economic and institutional conditions. Socialization is likely to occur and migrants accommodate to new practices which can have a substantial impact on their countries of origin.[2]

[1]Two other papers have addressed the role of migration on women's condition from a broader perspective. Diabate and Mesplé-Somps (2019) uses an original household-level database coupled with census data to analyze the extent to which Malian girls living in villages with high rates of return migrants are less prone to female genital mutilation. Tuccio and Wahba (2018), instead, applies a migration-induced transfer of gender norms to Jordan focusing on the role of return migration on three gendered dimensions: the self-perceived role of women in the Jordan society, females' freedom of mobility, and women's decision-making power.

[2]See Gordon (1964) for cultural assimilation.

An increasing number of papers have investigated the role of migration in diffusing norms and values of different natures. Spilimbergo (2009) has shown, for example, the role of foreign students in promoting democracy in their home country, provided that the foreign education has been acquired in democratic countries. Docquier *et al.* (2016) have shown that international migration is an important determinant of institutional quality, as measured by democracy and economic freedom. Focusing on returnees, Mercier (2016) recognizes the positive role of international migration on the quality of the leadership and the emergence of the elites. Analogously, a few micro studies have contributed to the literature on the transfer of political norms through migration. Chauvet and Mercier (2014) find a positive effect of Malian returnees from non-African destinations on origin country's participation rates and electoral competitiveness; Barsbai *et al.* (2017) show how Westward migrants contributed to overthrowing the Communist party in Moldova; and finally, Batista and Vicente (2011) show how migration to countries with better governance has increased the demand for political accountability in Cape Verde. The same mechanism has been also applied to attitudes toward fertility. Focusing on Egypt, Morocco, and Turkey, Fargues (2007) shows that fertility rates in sending countries are affected by the rates prevailing in their migrants' host countries. Beine *et al.* (2013) extend Fargues' conclusions providing evidence of a transfer of fertility norms from international migrants to 208 countries of origin. In a micro setting, Bertoli and Marchetta (2015) find that return migration to Egypt from other Arab countries characterized by higher fertility rates has had a significant and positive influence on the total number of children. In a similar vein, Daudin *et al.* (2018) show that internal migration between 1861 and 1911 in France contributes to convergence toward low birth rates by diffusing cultural and economic information about low-fertility behavior.

Our study belongs and contributes to this strand of literature raising the possibility that international migrants transmit back home through various channels attitudes toward gender parity.[3] This issue is explored by looking at the impact of international emigration on the share of female parliamentary participation between 1970 and 2010. The identification of the exact way through which migration affects attitudes toward women in the countries of origin is difficult both at macro and micro level unless suitable data are available and it goes beyond the scope of this paper. Our empirical analysis addresses whether a "transfer of norms" mechanism is in place and its causal direction. In particular, we show that total international migration to countries where the share of female parliamentary seats is higher increases source country female parliamentary participation in the lower chamber of the National Parliament.

There is strong anecdotal evidence on the role of international migration in shaping female political empowerment in the countries of origin. Correa (1998), for example, finds how the involvement of Puerto Rican female migrants in the New York political arena changed the social role of women as well as their husbands' viewpoint concerning their wives at the origin.[4] When Nydia Velasquez won the Puerto Rican

[3]Also the study by Neumayer and De Soysa (2011) focuses on attitudes toward gender parity but in a different environment. An analysis of *spatial dependence* puts forward the role of trade and FDI in fostering the empowerment of women. Specifically, it is suggested that the incentive to raise women's rights is stronger where, firstly, major trading partners and secondly, the major source countries for FDI themselves provide strong rights. Economic and social rights are taken from the Cingranelli and Richards' (2009) Human Rights Database, but there is no direct reference to political rights. Nonetheless, the role of other globalized outcomes such as migration has not been touched.

[4]"In Latin America men were always the leaders. Women in politics were seen as strange", p. 343.

https://doi.org/10.1017/dem.2020.7 Published online by Cambridge University Press

primary elections and ended up being elected as the first Puerto Rican congresswoman in the US, she was strongly supported by Latino voters willing to accommodate themselves to the idea of female leadership.

Pessar (2001) studied the behavior of Guatemalans emigrated to Mexico and then returned to their country. She finds that in 1995, on the occasion of a meeting with returnee leaders, Guatemalans were persuaded to sign a document affirming the desirability of making women and men equal owners of the land and the equal accessibility to the community governing directorates. Migrants also keep strong links with their family back home. In Africa, where migration of women is mainly circular, migrants do not break away permanently from their places of origin [Piore (1979)]. The 2011 Nobel Peace Prize Leymah Roberta Gbowee while visiting regularly her origin country, Liberia, where her children used to live, struggled for the safety of women and for women's right. She led the women's peace movement that brought an end to the Second Liberian Civil War in 2003 and contributed then to the election of Ellen Johnson Sirleaf, the first African female President.[5]

Migrants frequently contribute to the development of their village of origin through Home Town Associations (HTAs), which can offer women's empowerment projects.[6]

For example, the South Sudan Women's Empowerment Network (SSWEN), created by Sudanese United States-based migrants, has been deeply involved in building the new South Sudanese National State, whose independence dates back to the 9th of July 2011. The role of the Sudanese diaspora has been so relevant for the involvement of women in development programs (with particular emphasis on political decision-making), that Erickson and Faria (2011) describe diasporic Sudanese women as "new and increasingly important citizens and activists in the post-CPA (Comprehensive Peace Agreement) era".

Finally, external voting also helps in transferring new political values in countries of origin.[7]

In 1916, the province of British Columbia in Canada enabled military personnel overseas to vote in a referendum on women's suffrage which became effective then.[8]

The choice of female parliamentary participation is important for many reasons. Women constitute more than half of the global population. However, female electorate continues to be under-represented in economic and political decision-making bodies at all levels. According to 2010's Inter-Parliamentary Union

[5]Mrs Leymah Roberta Gbowee spent some time in Virginia where she received a Master's Degree in Peace Building at the Eastern Mennonite University (EMU). She has then resided in Ghana where she moved before the independence of Liberia.

[6]The HTAs are immigrant informal organizations, based in a common hometown, that bring members together for social, cultural, political empowerment, and economic development goals.

[7]External voting refers to the right that enables migrants to vote from abroad. Even if the constitutions of many countries guarantee the right to vote for everybody, voters who are outside their home country are often disenfranchised because of a lack of procedures enabling them to exercise the right to vote. According to the voting operations data from Ace (The Electoral Knowledge Network-http://aceproject.org/ace-en/topics/va/external-voting-a-world-survey-of-214-countries), voting outside the boundaries is not permitted for 27.8% countries against 50.6% cases in which citizens residing outside the country can vote and 21.6% cases under which voting is permitted under special conditions (being member of the armed forces, diplomatic staff, students, etc.). Moreover, even where admitted, external voting is associated with low participation rates and this can be due to security concerns, voter disinterest, difficult access to registration and voting facilities, and documentation issues.

[8]See "Voting from Abroad: The International IDEA Handbook", 2007.

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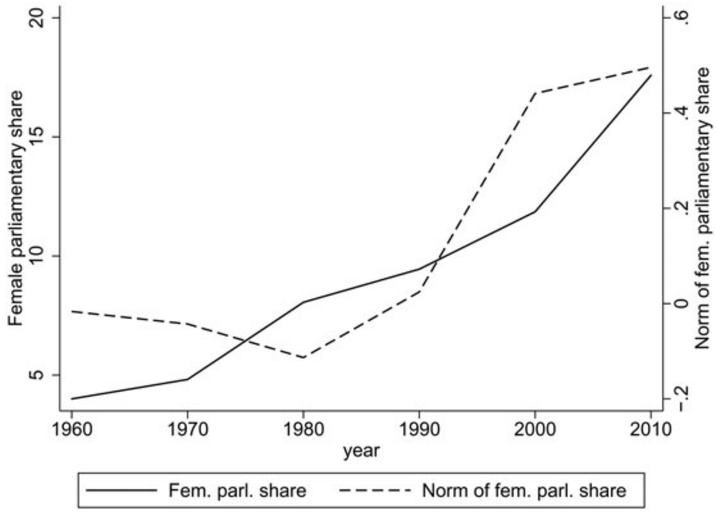

Figure 1. Female parliamentary share and norm of female parliamentary share over time (1960–2010).

(IPU) data, the international average representation of women in parliaments has increased over the years, but it is far short of gender parity (see Figure 1).

When revising the relationship between gender empowerment and economic development, Duflo (2012) states that there are two rationales for supporting active policies to promote women. The first is equity: women are currently worse-off than men, and this inequality between genders is unfair in its own right. The second regards the fundamental role women play for development. This is a central issue in policy makers' debates. Baskaran *et al.* (2018) use comprehensive data on competitive elections to India's state legislative assemblies, exploiting close elections between men and women to isolate the causal effect of legislator gender in a regression discontinuity design. They find evidence that women legislators are less likely to be criminal and corrupt, more efficacious, and less vulnerable to political opportunism.

Thomas (1991) shows that States in the US with a higher female representation have introduced and passed more priority bills dealing with the issues of women, children, and families compared to States with lower female representation. Funk and Gathmann (2008) examine survey data on all federal votes in Switzerland between 1981 and 2003, and find that female voters and politicians favor, among other things, public health provision, and equal gender rights. Besley and Case (2000) find that female legislators apply pressure to increase family assistance and to strengthen child support. When considering Indian data, Clots-Figueras (2011) finds that female legislators invest more than men in health, female teachers, early education, and favor "women-friendly" laws.

Along the same lines, Iyer *et al.* (2011) find that an increase in female representation in local government induces large and significant effects on reporting of crimes against women in India, thus favoring access to justice for women. On the theoretical side, De la Croix and Vander Donckt (2010) recognize the importance of female empowerment as a multidimensional concept which includes economic participation and opportunity, educational attainment, political empowerment, health, and survival. They argue that a

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range of socioeconomic outcomes are attached to gender equality, including improved children's development (through better health and education) and growth.

The rest of the paper is structured as follows. Section 2 describes the empirical model implemented to identify the impact of migration on the propagation of female political values at home. Section 3 deals with the datasets used to conduct the empirical analysis and provides some stylized facts. Section 4 goes through the main challenges to be addressed in the empirical analysis. Section 5 provides the empirical results and some robustness checks. Finally, Section 7 concludes.

2. The empirical model

To test for the impact of international migration on female parliamentary seats in the country of origin ($\text{seats}^F_{i,t}$) through a "transfer of norms" mechanism, we consider the following specification by decade[9]:

$$\begin{aligned}\text{seats}^F_{i,t} &= \alpha \text{seats}^F_{i,t-10} + \beta \sum_j \left[\frac{\text{mig}_{ij,t-10}}{\text{pop}_{i,t-10}} \times (\text{seats}^F_{j,t-10} - \text{seats}^F_{i,t-10}) \right] \\ &\quad + \sum_{m=1}^{n} \rho_m R_{i,t} + \mu_i + \varphi_t + \epsilon_{i,t}\end{aligned} \tag{1}$$

where:

- t refers to the year of interest and goes from 1970 to 2010; i refers to the country of origin and j to the country of destination.[10]
- $\text{seats}^F_{i,t}$ represents the female parliamentary share at time t in country of origin i.
- $\text{seats}^F_{i,t-10}$ represents the female parliamentary share at time $(t-10)$ in country of origin i.
- $\text{mig}_{ij,t-10}$ is the bilateral total migration stock from i to j at time $(t-10)$. The reason why we consider total migration instead of female migration is threefold. First, according to the message given on the occasion of the 100th International Women's Day by the Director of the Secretariat of the International Strategy for Disaster Reduction (UN/ISDR) Salvano Briceno, "Advancing gender perspectives and women's rights is not just a job for women, more men must advocate at a high level for the empowerment of women, and for the incorporation of gender budgeting into national and local development plans". Secondly, if we look at the gender composition of HTAs, there is no evidence that efforts to improve females' conditions are just pursued by female migrants. Recent developments have shown that policies and works towards gender equality face new challenges related to men's role and demands.[11] Thirdly, according to Doepke and Tertilt (2009), men care about the other gender in facing a trade-off between the rights they want for their own wives (namely none) and the rights of other women in the economy.
- $\text{pop}_{i,t-10}$ is the total population at time $(t-10)$ in country i.

[9]The time lag is equal to 10 years because migration data from Ozden *et al.* (2011) is only available by decade.

[10]See Appendix A for the list of countries in the sample.

[11]Website: http://www.womenlobby.org (Brussels, 7th October 2011).

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- $[(\text{mig}_{ij,t-10}/\text{pop}_{i,t-10}) \times (\text{seats}^F_{j,t-10} - \text{seats}^F_{i,t-10})]$ is the "norm" at time $(t-10)$ through which foreign female parliamentary participation is propagated in the country of origin. Unlike previous works, we multiply the migration rate component $\text{mig}_{ij,t-10}/\text{pop}_{i,t-10}$ with the difference between the parliamentary share at destination and that in the country of origin. We expect a positive effect if $\text{seats}^F_j > \text{seats}^F_i$. In other terms, the origin country takes advantage of the political environment at destination just if the female political conditions at destination are better than those at origin (we will have instead a "negative transfer of norm" if $\text{seats}^F_j < \text{seats}^F_i$ and no transfer if $\text{seats}^F_j = \text{seats}^F_i$). Moreover, the greater the difference, the stronger the effect.
- $R_{i,t}$ contains other traditional covariates of interest. We control for political exogenous variables such as the presence of a (*de jure*) democratic system in the country of origin i at time t; the occurrence of legal elections in between time t and time $t-10$ in country i and the number of legal elections with a proportional system between time t and time $t-10$ in country i. We account also for the female skill ratio in country i at time $t-10$ computed as the ratio of tertiary educated over illiterate females; the number of years since CEDAW (Convention on the Elimination of all Forms of Discrimination against Women) ratification at time t, GDP and trade norms at time $t-10$ in country i.
- μ_i and φ_t are country of origin and time fixed effects.

The main references are the studies of Spilimbergo (2009) and Beine *et al.* (2013). To determine the impact of students' migration on democracy at the origin, Spilimbergo (2009) regresses the index of democracy at time t in country i over the 5 years' lagged value of democracy in country i, the number of students abroad as a share of the total population in the sending country, the average level of democracy in the host countries, and the interaction between the two latter terms. The average level of democracy in the host countries is constructed as the weighted average of the institution in the host countries where the weights are given by the share of students from country i to country j over all students from country i. Beine *et al.* (2013) also apply the same specification in a cross-section setting to assess the impact of migration on source country fertility. The norm is constructed as the interaction between the (log of) fertility rate at destination with the size of the diaspora. With respect to previous studies, our norm differs in two aspects. First, it is able to control for asymmetries between the source country and destination's female political empowerment. Secondly, its weights are given by emigration rates in order to test whether the transmission of the norm depends on the intensity of migration. Spilimbergo (2009) and Beine *et al.* (2013) are prevented from doing it because of collinearity problems. The correlation between the norm, the migration rate, and the interaction term between the two is so high that they cannot infer anything on the intensity of migration. In Beine *et al.* (2013), in particular, this lack of significance is justified by the complexity of the transfer of norms' mechanism.[12]

3. Data and stylized facts

Our data set is a 10-year unbalanced panel spanning the period between 1970 and 2010, where the start of the date refers to the dependent variable (i.e., $t = 1970$, $t - 10 = 1960$).

[12]See Appendix C for a more detailed description of the differences with previous studies.

Countries enter the sample if women were eligible at time t and at time $t-10$. The country sample is selected on the basis of the availability of the data. We now describe the basic details of the data we use to measure our main variables of interest.

3.1 Political data

Political data on the proportion of seats held by women in national parliaments cover the time span 1960–2010 and relies on two different datasets.[13] Between 1960 and 2003, the database by Paxton *et al.* (2006) titled "Women in Parliament, 1945–2003: Cross-National Dataset" is used.[14] This data collection provides yearly information on women's inclusion in parliamentary bodies in 204 countries from 1945 to 2003. The dataset allows for the extensive, large-scale, cross-national investigation of the factors that explain women's attainment of political power over time and provides comprehensive international and historical information on women in a variety of political positions. Information is provided on female suffrage, the first female member of parliament, yearly percentages of women in parliaments (data refer to the percent of parliamentary women in the lower or single house of each country's national legislature), when women reached important representational milestones, such as 10%, 20%, and 30% of a legislature, and when women achieved highly-visible political positions, such as prime minister, president, or head of parliament. Political information for the remaining 7 years (from 2004 to 2010) has been taken from the World Development Indicators (WDI) 2014. Both Paxton *et al.* (2006) and WDI (2014) rely on IPU (www.ipu.org) data which make them compatible with each other.

The evolution of the world average female parliamentary share is described in Figure 1. Looking at the world average, the proportion of seats held by women in national parliaments has increased from about 4% in 1960 to about 18% in 2010. While the international average representation of women in parliaments has steadily increased, this is far short of gender parity. In 2010, 58 countries still have no more than 10% of female members in the parliament. Moreover, the gap in women parliamentary participation among regions is high, with the West Europe region being the most "feminized" with an average of 30% of women in the Parliament. The lowest shares belong to the Mena region, with an average of 11.1%.

Figure 2 represents cross-country differences in the average level of the proportion of seats held by women in national parliaments over the period 1960–2010. Among the countries with the highest proportions, we find European countries such are Nordic countries, but also some Latin American and Caribbean nations (Cuba at the top) and few African countries, such as Rwanda or Mozambique. Muslim countries and the Arab states lag well below the other countries.

In order to be consistent with migration data, we consider political data by decade for 1960, 1970, 1980, 1990, 2000, and 2010.[15]

[13]The exact definition of the variable is "Women in parliaments as the percentage of parliamentary seats in a single or lower chamber held by women".

[14]http://www.icpsr.umich.edu.

[15]Preferring 10 years data to yearly data is important for at least three reasons. It avoids migration and human capital data interpolation; the persistence due to political legislatures is reduced and a longer period for the occurrence of a "transfer of norm" mechanism is taken into account. It might be indeed the case that migrants require more than one year before integrating and then transmit new values in their home country.

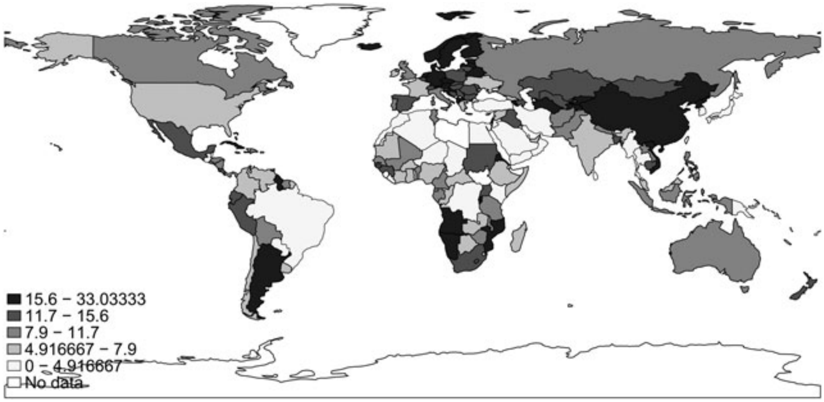

Figure 2. Cross-country differences in female parliamentary share, 1960–2010 average.

In the political database by decade, there are some missing values due to the absence of the parliament (i.e., coup d'etat, dictatorship war, "false elections", no sovereignty, or colonialism).[16] For the construction of the norm and the empirical analysis, we consider the reduced sample (female parliamentary share is missing because the Parliament is absent). In addition, countries enter the sample if women were eligible at time t and at time $t - 10$.[17]

3.2 Migration data

Migration information relies on the recently released bilateral database described in Ozden *et al.* (2011).[18] They provide bilateral migration stocks disaggregated by gender by decades (for the period 1960–2000) for 232 countries, relying primarily on the foreign-born concept. Over 1,000 census and population register records are combined to construct decennial matrices corresponding to census rounds for the entire period. In doing so, the authors provide for the first time, a complete picture

[16]In the original Paxton *et al.* (2006) political database, there are indeed three types of missing values. The so-called "true missing" due to the fact that the authors could not find positive data, a second type of missing due to coup d'etat, and a third type of missing due to the absence of the Parliament. The absence of the parliament can be due in turn to several factors: the presence of a dictatorship, "false elections" or the absence of sovereignty, i.e., colonial reasons. Concerning true missing values, we have transformed them in an historical/political compatible way into either missing values, zeroes, or positive observed values using an additional political dataset from Armingeon and Careja (2008) as explained in Appendix D.

[17]The probability for a woman to "be eligible" is equal to 0 if the Parliament is absent or if women do not have the right to be voted yet. In the first case, we do not observe data on female parliamentary share. In the second case, female parliamentary share is 0 because there are parliamentary sessions, but women are not eligible (this happens for 21 observations which are excluded from our sample). In the case of eligibility, we can have two conditions: the female parliamentary share is positive or the female parliamentary share equals zero because women have the right to stand for office but nobody vote for them. Actually, the female parliamentary share can also be zero when there are parliamentary sessions but women do not run for any political position. Since we do not have data on female political entry, we assume that there are some women who run for the position in any case. Focusing on the reduced sample, we disregard selection issues due to female political eligibility.

[18]http://data.worldbank.org/data-catalog/global-bilateral-migration-database

of bilateral global migration over the second half of the twentieth century, including for the first time also South-South migration. According to Ozden *et al.* (2011), total migration has steadily increased over the period 1960–2000. The data reveal that the global migrant stock increased from 92 million in 1960 to 165 million in 2000. This figure is largely driven by migrants from developing countries. In section 5.7, for robustness, migration data are complemented by the UN migration dataset for the period 2000–2010. Altogether, these data reveal that the largest increase pertains to the share of immigrants from non-OECD to OECD countries (from 16% in 1960 to 34% in 2010), while the lowest share belongs to migration from OECD to non-OECD countries (from 11% in 1960 to 2% in 2010).

When combining migration with political data, the geographical dimension of the initial migration dataset has been complemented, reconstructing data for Czechoslovakia, the Socialist Federal Republic of Yugoslavia, and USSR for which data were available following their political split.[19]

Figure 3 represents cross-country differences in the average level of the norm of female parliamentary share over the period 1960–2010. A large heterogeneity both within developing and developed countries is observed.

The evolution of the world average norm is described in Figure 1. Looking at the world average, the norm has increased from about −0.02 in 1960 to about 0.5 in 2010. This increase might be due to an increase in international migration, but also in the average representation of women in parliaments. Figure 1 shows that the world average female parliamentary share and the average norm are positively correlated.[20]

3.3 Other data

Additional explanatory variables have been collected using the following databases.

Data on total population is provided by the World Population Prospects, the 2012 Revision, by the United Nations (2013). GDP per capita comes from the World Bank Development Indicators 2014. Female human capital indicators used to construct the female skill ratio are taken from Barro and Lee (2013). Barro and Lee' s data are available every 5 years. The indicator for democracy is the Polity2 indicator from the POLITY IV data set. This is a combined score that reflects several aspects such as the presence of institutions and procedures through which citizens can express effective preferences about alternative policies and leaders; the existence of institutionalized constraints on the exercise of power by the executive power; and the guarantee of civil liberties to all citizens in their daily lives and in acts of political participation. It ranges from −10 to +10 and it can be considered as a *de jure* indicator of the quality of institutions in the country.

Data on legal elections and electoral systems (proportional, majoritarian, mixed, and multi-tier) are from Golder (2005). Data on CEDAW ratification has been collected by us. We construct a variable which indicates the number of years since the convention has been ratified by the country.[21] Data about religion which identifies countries

[19]See Appendix B for the detailed reconstruction of these cells.

[20]The female parliamentary share and the norm exhibit a correlation of 0.850 over the period 1960–2010.

[21]The Convention was opened for signature at the United Nations Headquarters on 1 March 1980. Although the United States never ratified the convention, CEDAW has become the main international legal document on women's rights.

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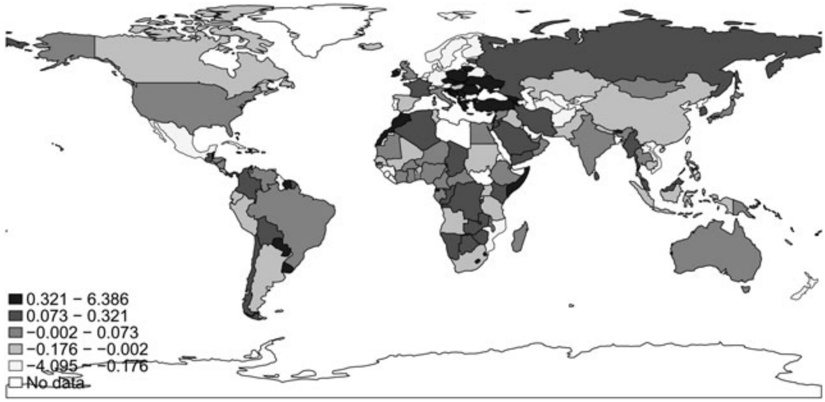

Figure 3. Cross-country differences in the norm of female parliamentary share, 1960–2010 average.

with more than 80% of the Muslim population are taken from La Porta *et al.* (1999). Data on trade are taken from Feenstra *et al.* (2004) who provide yearly world trade flows in the time span 1962–2000.

4. Econometric issues

OLS regressions generate inconsistent estimates in the presence of omitted variables, reverse causality, reflection problems, and other endogeneity issues. Then, we first estimate equation (1) using fixed-effect cross-country regressions, which account for omitted variables that do not largely vary over time. In order to account for other endogeneity issues, we instrument the main regressor of interest with an external instrument in a standard 2SLS. Finally, as we consider a standard dynamic panel specification, we use SYS-GMM dynamic panel estimations using internal instruments in order to account for persistency and endogeneity of all the covariates.

4.1 Omitted variables

As using pooled OLS, we will not control for possible mis-specifications due to unobserved characteristics, which may jointly affect international migration and the share of female parliamentary seats, we first estimate our empirical model by introducing country fixed effects. Although country fixed effects cannot capture determinants that are both country- and time-specific, they account for many unobservable characteristics. However, it should be noticed that other factors which may affect female parliamentary seats, such as female education, or the quality of political institutions, are very persistent. Therefore, the inclusion of country fixed effects in the regression model mostly accounts for them.

4.2 Reverse causality

A key issue when using fixed-effect regressions to study the relationship between international migration and political environment is the endogeneity of "the norm" $\left(\sum_j (\text{mig}_{ij,t-10}/\text{pop}_{i,t-10}) \times (\text{seats}^F_{j,t-10} - \text{seats}^F_{i,t-10})\right)$ due to a reverse causality issue.

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From one side, gender (in) egalitarian practices can act as a (push) pull factor: international migrants move to countries with better prospects for women [Ferrant and Tuccio (2015), Nejad and Young (2014)] or international migration acts as a way out of political discrimination [Ruyssen and Salomone (2018), Baudassé and Bazillier (2014), Hugo (2000)]. From the other side, gender inequality might prevent intended migrants from realizing their aspirations [Ruyssen and Salomone (2018)], as it is likely to be correlated with poverty and institutional quality (bad institutions, or low government effectiveness, can be responsible for large administrative costs).

This means that a downward or upward bias can be driven by reverse causality. An external instrumentation strategy is required in order for the coefficient of the "norm" to be unbiased.

The same argument holds for reflection issues [Manski (1993)]. The "norm" can be endogenous because if the equations for each country i were written in a system, the female parliamentary share would appear either as regressand for country i and as regressor within the norm for country $i+1$, $i+2$, etc.

We instrument the "'migration" part of the norm using a gravity-type equation model based on exogenous geographical and cultural bilateral variables (while controlling for origin and destination fixed effects). We predict bilateral migration stocks, in line with gravity-type equation used to predict trade bilateral pairs [e.g., Frankel and Romer (1999)].[22] As in Feyrer (2019), who builds a time-varying geographic instrument for trade based on a gravity-type equation, distances are interacted with time dummies. This introduced time variation captures common shocks in the changes in transportation technology over time which occur at the global level (distances are shorter and shorter time goes by: think about changes in air travel transportation costs). As long as changes in transportation technologies are common to all countries, these time series changes will be exogenous with respect to any particular country, but they will have different effects across country pairs, depending on the relative geographic position. While interacted distances with time dummies can be seen as a very good proxy of the air distance between two countries, the movement of people can occur also through other ways such as the sea. For this reason, we control also for sea distances (interacted with time dummies), which indicate the sea distance between the relevant ports of each one of the country pairs. The following gravity model is estimated:

$$\begin{aligned}\text{mig}_{ij,t} = \alpha + \delta_{1,t}\log(\text{dist})_{ij} + \delta_{2,t}\log(\text{seadist})_{ij}\\ + \beta_1\text{commonborder}_{ij} + \beta_2\text{language}_{ij} + \gamma_i + \gamma_j + \gamma_t + \epsilon_{ijt}\end{aligned}$$

where the dependent variable $\text{mig}_{ij,t}$ is the stock of immigrants from country i to country j at time t. The explanatory variables from the CEPII database are the distance between country pairs (interacted with time dummies), a dummy for whether country i and j share a common border, a dummy for speaking a common language. Bilateral sea distances (also interacted with time dummies) are taken from Bertoli *et al.* (2016). We include origin and destination fixed effects, which absorb the origin-specific and the destination-specific regressors, and time fixed effects.

[22]This method is standard in the migration literature. [See also Beine et al. (2013), Ortega and Peri (2014), Alesina et al. (2016), Docquier *et al.* (2016).]

The presence of a large number of zeroes in bilateral migration stocks gives rise to econometric concerns about possible inconsistent OLS estimates. We estimate the above model using the Poisson regression by pseudo-maximum likelihood. We use the PPML command in Stata which implements the method of Santos Silva and Tenreyro (2011) to identify and drop regressors that may cause the non-existence of the (pseudo-) maximum likelihood estimates. Standard errors are robust and clustered by country pairs.

The resulting exogenous norm to be used as an external instrument is:

$$\sum\nolimits_j \left[\frac{\widehat{\mathrm{mig}_{ij,t-10}}}{\mathrm{pop}_{i,t-10}} \times (\mathrm{seats}^F_{j,t-10} - \mathrm{seats}^F_{i,t-10}) \right].$$

4.3 Endogeneity of other regressors

Although the instrumental variable strategy previously described corrects for the "migration" endogeneity part of the "norm", it does not account for the endogeneity of other regressors. For example, female human capital can be an important determinant for female political participation; however, it could be that women in parliament affect the incentives of women to acquire education. In addition, female parliamentary share is very persistent. A standard panel dynamic framework (with the female parliamentary participation at time t regressed over its 10 years lag) is required [equation (1)]. The introduction of the lagged dependent variable can induce potential biases in the estimation. In addition, the norm is correlated with the lagged dependent as this enters our definition of the norm. In order to overcome endogeneity issues due to the lagged dependent and other lagged explanatory variables, we consider an SYS-GMM technique.

The system GMM estimator accounts for unobservable heterogeneity and it is preferable to a standard fixed-effects estimator since the inclusion of the lagged dependent variable in a fixed-effects model would lead to the so-called Nickell (1981) bias because the lagged dependent variable is correlated with the error term. However, it should be noticed that in our context, this bias should not be very sizable, given the large time span of our analysis. However, the fixed-effect estimator is not recommended when data are very persistent as it exacerbates measurement error bias [Hauk and Wacziarg (2009)], whereas the system GMM is the most appropriate estimator when time series are very persistent as in our case [see Bond *et al.* (2001)].

The system GMM estimator combines the regression in differences with the regression in levels in a single system. The instruments used in the first differentiated equation are the same as in Arellano and Bond (1991), but the instruments for the equation in level are the lagged differences of the corresponding variables. In order to use these additional instruments, a moment condition for the level equation, which implies that first differences of pre-determined explanatory variables are orthogonal to the country fixed effects, must be satisfied.

We test the validity of moments conditions by using the test of overidentifying restrictions proposed by Hansen and by testing the null hypothesis that the error term is not second order serially correlated. Furthermore, we test the validity of the additional moment conditions associated with the level equation using the Hansen difference test for all GMM instruments.

A particular concern related to this method is the risk of instrument proliferation. Using too many instruments can bias the GMM estimation results and weaken the Hansen test of the instruments' joint validity [Roodman (2009)]. We have, therefore, kept the number of instruments lower than the number of groups [as Roodman (2009) suggests].

The SYS-GMM estimator provides consistent and unbiased estimates but it depends on the particular set of instruments used. However, it is recognized that, with very conservative data, it is the best available estimator [Blundell and Bond (1998), Arellano and Bover (1995)].

Finally, a sensitivity analysis will also be conducted to check the robustness of the results to the exclusion of certain countries (e.g., socialist countries, Sub-Saharan African countries, and Muslim countries) whose characteristics may exacerbate reverse causality problems.

5. Estimation results

The results are organized in sub-sections. We first provide FE results. Secondly, a two-step procedure is implemented in order to account for endogeneity bias. Subsection 5.2 presents the pseudo-gravity type results, while subsection 5.3 presents 2SLS-FE results using the external instrument. Thirdly, in subsection 5.4, we estimate the dynamic specification as in equation (1) using the SYS-GMM technique, using internal instruments. Fourth, we conduct a sensitivity analysis to check the robustness of our results to the exclusion of certain groups of countries (socialist countries, Sub-Saharan African countries, and Muslim countries). Fifth, we consider an alternative definition of the norm to test for immigration-induced transfers of norms on gender political empowerment. Sixth, we test the robustness of our results considering annual data in order to take into account the complete political evolution of each country.

5.1 Panel analysis with FE results

Table 1 reports FE estimates from equation (1). No additional controls other than the lagged dependent and the lagged index of female parliamentary share (i.e., the "norm") are considered. Standard errors are robust and clustered by country of origin. We observe that in columns (1), the estimated coefficient of the norm is positive and statistically significant at the 10% level.

For comparison with previous studies, column (2) reports FE results for the base model without additional controls following the Spilimbergo (2009)'s specification, which includes the total emigration rate, an index for institutional quality in host countries, and an interaction between the two terms.[23] The new index of female parliamentary share is positive and statistically significant. The total emigration rate (which measures the direct effect of migration) is not statistically significant as in Spilimbergo (2009), as well as the interaction term, which captures the intensity of emigration with respect to the origin population in the transfer of female political values. Following Spilimbergo's specification, the lagged migration rate has no impact on female parliamentary seats at home while the quality of political institutions in host countries has a strong impact on political institutions at home. It is unclear whether this effect increases with the number of migrants abroad [so as in the original work by Spilimbergo (2009) and Beine *et al.* (2013)]. As the index of female

[23]See Appendix C for a detailed description of the Spilimbergo (2009)'s set up.

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Table 1. Estimations: fixed effect

	(1) Fem. Parl. share	(2) Fem. Parl. share
Female parliamentary share (lagged)	0.4777***	0.6147***
	(0.072)	(0.11)
Norm of fem. parl. share (lagged)	1.115*	
	(0.669)	
Total migration rate (lagged)		7.263
		(7.46)
Norm of fem parl. share *à la Spil.* (lagged)		0.2642**
		(0.117)
Interaction term *à la Spil.* (lagged)		0.1976
		(1.09)
Constant	12.22***	9.816***
	(1.11)	(1.68)
Country fixed effects	Yes	Yes
Year fixed effects	Yes	Yes
Observations	551	551
Number of countries	169	169
R^2	0.5093	0.5109

*Significant at the 10% level; **5% level; ***1% level. Robust standard errors clustered by country in parentheses.

parliamentary share and the interaction term are highly correlated, low significance levels could be due to collinearity problems. Our alternative specification directly captures the intensity of migration in transferring the norm, thus overcoming this concern. We can therefore say that the quality of female political participation in host countries has a positive correlation with female parliamentary share in the country of origin, which increases with the share of migrants abroad with respect to the total population in the country of origin.

5.2 Instrumentation for the norm

Table 2 presents estimation results from the gravity model used to predict the bilateral exogenous migration component of the "norm". Not surprisingly, geographic characteristics are strong determinants of bilateral migration stocks. As proxies of migration costs, contiguity, and linguistic links favor migration, geographical and sea distances are negatively correlated to those stocks.[24]

[24]The estimated coefficients of the interaction of sea distance with time dummies and of the interaction of air distance with time dummies are jointly statistically different from 0. Wald tests on the joint significance of the estimated coefficients of sea distance and of air distance are, respectively: $\chi^2_5(\text{seadist}) = 27.70$, $P - \text{value}_{\text{seadist}} = 0.00$; $\chi^2_5(\text{airdist}) = 58.95$, $P - \text{value}_{\text{airdist}} = 0.00$. Note that we

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5.3 Panel analysis with 2SLS

In Tables 3 and 4, we correct for endogeneity using 2SLS regressions with country and time fixed effects.

Columns (1) and (2) of Table 3 show both the first and second stages when instrumenting the norm with $\sum_j [(\widehat{\text{mig}_{ij,t-10}}/\text{pop}_{i,t-10}) \times (\text{seats}^F_{j,t-10} - \text{seats}^F_{i,t-10})]$. The first stage shows that our instrument is positive and statistically correlated with the norm. Since the Kleibergen–Paap rk Wald F statistic is well above the critical values reported by Stock and Yogo (2005), we reject the hypothesis that our instrument is weak.

The second stage shows that the estimated coefficient of the "norm" is positive and statistically significant.

Compared with OLS in Table 1, the 2SLS coefficient is larger, suggesting that the OLS coefficient might suffer from a reverse causality (downward) bias: emigration decreases when women's representation in parliament is higher.

Table 4 adds traditional political and non-political covariates to the above specification.

In column (1), we add a measure of democracy as a proxy for the quality of institutions in the country.

Column (2) refers to the model which we will consider as our baseline specification henceforth. It contains the lagged dependent, the lagged "norm" of female parliamentary share, a lagged measure of female human capital, and a measure of *de jure* democracy.

Female human capital can be important in explaining female political empowerment: women need human and financial capital (gained through education and work experience) to stand for office [Paxton and Kunovich (2003)].[25] As a proxy for female human capital, we generate the ratio between the number of females aged more than 25 years old with tertiary completed education and females with no schooling. In 2SLS regressions with country fixed effects, the estimated coefficient is positive and generally statistically significant.

An indicator of democracy is also considered. Indicators of democracy measure the general openness of political institutions and combine several aspects such as the presence of institutions and procedures through which citizens can express effective preferences about alternative policies and leaders; the existence of institutionalized constraints to the exercise of power by the executive power; and the guarantee of civil liberties to all citizens in their daily lives and in acts of political participation. In our case, we consider a composite index called Polity2 that ranges from −10 to +10, with 10 corresponding to the most democratic set of institutions. It is worth reminding that Polity2 captures the quality of "*de jure*" institutions, and it is not based on perceptions (not capturing the quality of *de facto* institutions). The effect of democracy on women's political representation may be ambiguous. On the one hand, it may be easy for women to be elected to a powerless parliament or under an authoritarian system, built on egalitarian ideologies such as ex-communist countries

cannot reject the null hypothesis of the equality of the interacted coefficients for air distance and sea distance, respectively.

[25]Women's workforce participation may also favor women's political participation. We indirectly control for workforce participation through human capital (of course, the two variables are highly correlated). We cannot introduce female labor force participation rate as a control variable, as data are available only from the 80s.

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Table 2. Gravity regressions (dep = bilateral migration stocks)

	PPML
Common border	1.278***
	(0.175)
Common official language	1.308***
	(0.176)
Log (Sea Dist) × 1960	−0.277*
	(0.164)
Log (Sea Dist) × 1970	−0.328**
	(0.138)
Log (Sea Dist) × 1980	−0.231*
	(0.123)
Log (Sea Dist) × 1990	−0.286**
	(0.115)
Log (Sea Dist) × 2000	−0.334***
	(0.0776)
Log (Dist) × 1960	−0.932***
	(0.170)
Log (Dist) × 1970	−0.888***
	(0.154)
Log (Dist) × 1980	−0.898***
	(0.142)
Log (Dist) × 1990	−0.778***
	(0.136)
Log (Dist) × 2000	−0.718***
	(0.113)
Constant	14.62***
	(1.032)
Origin fixed effects	Yes
Dest. fixed effects	Yes
Year fixed effects	Yes
Observations	141,486
R^2	0.581

*Significant at the 10% level; **5% level; ***1% level. Robust standard errors clustered by country pairs in parentheses.

where female parliamentary participation was high. On the other hand, more democratic countries may favor women's political participation. In 2SLS regressions with fixed effects, Polity2 is negative and statistically significant.

Table 3. Estimations 2SLS-FE

	(1) First stage Norm of fem. parliamentary share (lagged)	(2) Second stage Fem. parliamentary share
Female parliamentary share (lagged)	0.0168**	0.4979***
	(0.008)	(0.071)
Norm of fem. parl. share (lagged)		1.483**
		(0.686)
Pre. Norm of fem. parl. share (lagged)	1.395***	
	(0.163)	
Year fixed effects	Yes	Yes
Country fixed effects	Yes	Yes
Observations	517	517
Number of countries	135	135
Kleibergen–Paap rk Wald *F* statistic		73.08
10% maximal IV size		16.38

*Significant at the 10% level; **5% level; ***1% level. Robust standard errors clustered by country in parentheses. Kleibergen–Paap rk Wald *F* statistics to be compared with the Stock–Yogo critical values for weak instrumentation.

In column (3), we control for legal election, which indicates the number of elections to national lower chamber occurred between time t and $t-10$, and for the electoral system being proportional. It is recognized in the political science literature that proportional systems, rather than majority ones, help women to access the political system [e.g., Paxton et al. (2010), Jalalzai and Krook (2010)]. Proportional systems make use of multi-member districts, which implies that more than one candidate can be elected from a particular district, and often have closed party lists, which means that citizens vote for the party lists of candidates rather than individual candidates. Under a list system, parties may feel compelled to nominate women in order to balance the list. Moreover, the higher the district magnitude, the greater the probability for a woman to be nominated, if the political party is expecting to win several seats in the district.[26] The election variable is generally positive, but not significant as well as the proportional nature of the electoral system.

[26]Of course, concerning electoral system's characteristics, the introduction of an electoral gender quota may encourage greater representation of women. Unfortunately, we cannot directly control for quotas, as data are available only for the most recent election years (see the Global Database of Quotas for Women at http://www.quotaproject.org/ by the International Institute for Democracy and Electoral Assistance-IDEA). The rapid diffusion of gender quota across countries has indeed occurred within the last 15 years. The inclusion of time dummies in our specification should capture the general increase in female representation due to the contemporaneous introduction of gender quotas in political systems. In addition, as many studies find that the greatest impact of quotas occurs under electoral systems with closed list and higher district magnitude [see Jalalzai and Krook (2010)], controlling for the

Table 4. Estimations 2SLS-FE

	(1) Female parl. share	(2) Female parl. share	(3) Female parl. share	(4) Female parl. share	(5) Female parl. share	(6) Female parl. share	(7) Female parl. share
Female parl. share (lagged)	0.4888***	0.472***	0.4519***	0.4735***	0.4829***	0.4018***	0.2671**
	(0.075)	(0.081)	(0.102)	(0.084)	(0.097)	(0.136)	(0.121)
Norm of fem. parl. share (lagged)	2.269***	2.167***	2.8***	2.162***	2.478***	1.644***	1.932**
	(0.734)	(0.637)	(0.82)	(0.627)	(0.695)	(0.604)	(0.833)
Democracy index	−0.2913**	−0.279**	−0.2171***	−0.2752**	−0.1	−0.3106**	−0.2594***
	(0.121)	(0.129)	(0.08)	(0.134)	(0.121)	(0.146)	(0.099)
Skill ratio for females (lagged)		0.0191**	0.0196*	0.0191**	0.0172	0.0163	0.0093
		(9.4×10^{-3})	(0.011)	(9.5×10^{-3})	(0.011)	(0.012)	(0.013)
Proportional electoral system			0.1128				0.8798**
			(0.447)				(0.449)
Legal election (sum)			0.4078				−0.6139
			(0.438)				(0.622)
Years since CEDAW ratification				−0.0308			−0.0571
				(0.164)			(0.118)

Norm in GDP per capita (lagged)					-5.1×10^{-5}		8.3×10^{-4}**
					(2.6×10^{-4})		(3.7×10^{-4})
Norm in trading partner (lagged)						0.0461	−0.1438
						(0.17)	(0.117)
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Country fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	466	417	250	417	378	328	197
Number of countries	119	103	65	103	98	93	60
Kleibergen–Paap rk Wald *F* statistic	38.39	52.89	99.92	52.94	66.4	82.35	165.4

*Significant at the 10% level; **5% level; ***1% level. Robust standard errors clustered by country in parentheses. Kleibergen–Paap rk Wald *F* statistics to be compared with the Stock–Yogo critical values for weak instrumentation.

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From columns (4)–(6), other potential traditional explanatory variables are added to the baseline specification. In particular, we consider years since CEDAW ratification, a norm based on GDP per capita, and a trade "norm" variable.

The CEDAW ratification implies that countries, which ratified the convention, should meet the minimum standards to reach equal women rights. Moreover, the countries involved should regularly provide the measures they have taken to reach this goal. As in True and Mintrom (2001), CEDAW is not significant (column 4) and this can be due to two reasons. First, some countries decide to ratify just because of international visibility with little intention to change gender relation; secondly, some countries ratify later as a consequence of lack of bureaucratic conditions.

In column (5), we control for the risk that the positive and significant effect of the migration norm could be driven by the fact that what matters is only migration toward richer countries. An alternative norm, based on bilateral migration weights and the difference between GDP per capita at destination and origin, has been constructed. This alternative norm is negative and not statistically significant. It should also be noted that, from one side, development itself matters for women [Burn (2005)][27], and from the other side, the correlation between GDP per capita and women's presence in parliament is not clear cut, depending on whether women have effective power in the country. For instance, in Central America, where quota laws are less common, there exists a positive and strong correlation between GDP per capita and female politicians in parliament. In South America, where quotas predominate, the opposite occurs: GDP per capita and women in parliament are negatively correlated [Mala and Piscopo (2010)]. In addition, the estimated coefficient is biased due to reverse causality (women in power promote development).

Column (6) considers a trade index, because economic integration may also convey cultural norms, supporting women's political participation. The trade norm is built in a symmetrical way to the migration one, constructing a weighted average of the difference in female parliamentary share with trading partners where the weights are given by the share of trade between the country of origin and the trading partner over total trade (e.g., $\sum_j [(\text{trade}_{ij,t-1}/\text{Trade}_{i,t-1}) \times (\text{seats}^F_{j,t-1} - \text{seats}^F_{i,t-1})]$). Our estimated coefficient is positive, meaning that trade is a measure of openness that goes in the same direction to migration, and not statistically significant.[28]

In the last column, we include all the explanatory variables. Our main results do not change again.

The magnitude of the estimated coefficients of the norm (i.e., the short-run effect) is always positive and statistically significant and it varies between 2.8 and 1.644. Focusing on our baseline specification (column 2), the estimated coefficient of the norm is 2.167. This means, for example, that a one standard deviation increase in the norm raises female parliamentary share by 1.39 percentage points.[29] This is a reasonable

proportional nature of electoral system means also indirectly controlling for the implementation of gender quotas.

[27]As explained in Bertocchi (2011), the logarithm of GDP per capita can be also considered as a proxy for the gender wage gap, given the strong negative correlation between the two measures.

[28]This result has to be taken with caution. We have indeed to notice that the trade dataset has much more missing values than the migration dataset. In unreported robustness checks, we control for the general openness of the country, using trade/gdp (lagged) as control variable. The openness indicator is not statistically significant.

[29]The effect is computed as follows: 2.167*(0.64) = 1.39

magnitude relative to the mean of female parliamentary share in the baseline sample (12.8), implying an increase of about 11%.[30]

It should also be noticed that most of the control variables are highly persistent; therefore, their effects can be captured by the inclusion of country fixed effects.

Finally, the bottom of the tables reports information regarding the performance of the external instrument we are using. For each specification, we report the Kleibergen–Papp *F* test (KP), which allows us to reject the null hypothesis of weak instrumentation.

5.4 Panel analysis with SYS-GMM

Table 5 reports estimation results in system GMM (the most appropriate estimator in this context). Our dependent variable is the share of seats held by women in the lower or single house in the National Parliament. In particular, column (1) refers to our baseline specification which contains the lagged female parliamentary share, the lagged norm of female parliamentary share, a measure of female human capital, and a measure of democracy. Columns (2) to (5) add additional controls as in the previous sub-section. Standard errors are robust and clustered by country group.

The coefficient for the lagged dependent is positive and statistically significant, and it ranges between 0.81 and 0.88.

In general, results are fairly similar to 2SLS-FE estimations. With respect to previous regressions, the estimated coefficient of Polity2 is no more statistically significant. The proportional electoral system is now positive and statistically significant.

In addition, endogenous variables can be instrumented using their own lagged values.

Importantly, in all the specifications, the norm remains positive and statistically significant, implying that total emigration is a positive and important channel through which female parliamentary share in the origin country raises. The magnitude of the estimated coefficients (i.e., the short-run effect) varies between 1.802 and 1.245. If we consider our baseline specification, the estimated coefficient equals to 1.728. This means that a one standard deviation increase in the norm raises female parliamentary share by 1.11 percentage points.[31] This is a reasonable magnitude relative to the mean of female parliamentary share in the baseline sample (12.8), implying an increase of about 8.7%.[32,33]

[30]The implied long-run effects are large and vary between 2.8/(1−0.4519) = 5.11 and 1.644/(1−0.402) = 2.75. In the baseline specification, the long-run estimated coefficient is 4.10, implying that a one standard deviation increase in the norm raises female parliamentary share by 2.63 percentage points in the long-run. In col. 2, given the estimated coefficient of the lagged dependent, this means that it takes about 13 years (1.31 periods of 10 years) to close half of the gap with the long-run level of female parliamentary share when a shock occurs. Note however that these point estimates are not the most preferred ones in our empirical analysis as the estimated coefficients of the lagged dependent in fixed-effects model are biased. See section 4.3.

[31]The effects are computed as follows: 1.728*(0.64) = 1.11.

[32]The implied long-run effects vary between 1.802/(1−0.853) = 12.26 and 1.245/(1−0.885) = 10.7. In our baseline specification, the long-run estimated coefficient equals to 1.728/(1−0.875) = 13.8. This means that a one standard deviation increase in the norm raises female parliamentary share by 13.8*0.64 = 8.8 percentage points in the long-run. These effects are relevant, however consider that given the estimated coefficient of the lagged dependent, this means that it takes 55 years (5.55 periods of 10 years) to close half of the gap with the long-run level of female parliamentary share when a shock occurs.

[33]In order to assess the importance of these effects at country-specific level, in section 6 we simulate the counterfactual female parliamentary share obtained in two extreme cases.

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Table 5. Estimations with SYS-GMM

	(1) Female parl. share	(2) Female parl. share	(3) Female parl. share	(4) Female parl. share	(5) Female parl. share	(6) Female parl. share
Female parl. share (lagged)	0.8751***	0.8094***	0.8772***	0.8533***	0.8845***	0.8049***
	(0.096)	(0.129)	(0.097)	(0.106)	(0.166)	(0.199)
Norm of fem. parl. share (lagged)	1.728***	1.533**	1.717***	1.802***	1.245***	1.663**
	(0.559)	(0.588)	(0.56)	(0.588)	(0.409)	(0.716)
Democracy index	−0.0049	0.074	−0.0015	0.0357	−0.0048	0.1556
	(0.054)	(0.071)	(0.057)	(0.057)	(0.064)	(0.115)
Skill ratio for females (lagged)	0.0142*	0.0137*	0.0141*	0.0144*	0.0146*	0.0137*
	(8.4×10^{-3})	(8.1×10^{-3})	(8.4×10^{-3})	(8.3×10^{-3})	(8.0×10^{-3})	(8.2×10^{-3})
Proportional electoral system		0.8274***				0.9474**
		(0.292)				(0.369)
Legal election (sum)		0.0265				−0.377
		(0.367)				(0.532)
Years since CEDAW ratification			−0.0332			−0.1902
			(0.092)			(0.132)
Norm in GDP per capita (lagged)				3.1×10^{-4}		1.3×10^{-4}

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				(2.7×10^{-4})		(4.2×10^{-4})
Norm in trading partner (lagged)					0.0429	−0.0674
					(0.186)	(0.176)
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Country fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	438	275	438	403	351	220
Number of countries	124	90	124	123	116	83
Number of Instruments	48	50	49	62	53	69
Arellano–Bond test for AR (1)	7.9×10^{-5}	9.6×10^{-4}	8.7×10^{-5}	1.9×10^{-5}	3.6×10^{-6}	0.0083
Arellano–Bond test for AR (2)	0.3207	0.2765	0.3186	0.534	0.1692	0.4484
Hansen test *p*-value	0.2007	0.1791	0.1732	0.5628	0.1876	0.3468
Difference Hansen test *p*-value	0.543	0.333	0.468	0.568	0.500	0.340

*Significant at the 10% level; **5% level; *** 1% level. Robust standard errors clustered by country in parentheses. SYS-GMM estimations.

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In all the empirical specifications just mentioned, we also take into account country and time fixed effects; therefore, the results are robust to all country-specific time-invariant characteristics which may influence female political representation. They encounter religion, colonial history, and many other unobservable characteristics.[34,35]

In the SYS-GMM estimations, the instruments used in the first differentiated equation are the same as in Arellano and Bond (1991), but the instruments for the equation in level are the lagged differences of the corresponding variables.[36] In our specifications, the lagged dependent variable is instrumented using from its own first to fifth lags. Our variables of interest, i.e., the lagged index of female parliamentary share, the lagged female human capital, the lagged index in trading partners, the lagged index of GDP per capita, are instrumented using their own first to fifth lags.

The legal election variable, the proportional system variable, the democracy indicator, and the CEDAW variable are considered as exogenous.

We test the validity of moment conditions by using the test of overidentifying restrictions proposed by Hansen and by testing the null hypothesis that the error term is not second order serially correlated. Furthermore, we test the validity of the additional moment conditions associated with the level equation using the Hansen difference test for all GMM instruments. The tests confirm the validity of our instruments.[37]

5.5 Robustness tests: heterogeneity in sample

In order to test for the robustness of our empirical results, we estimate our baseline specification in selected sub-samples. First of all, we exclude socialist countries (i.e., countries which belonged to the Iron Curtain). In the former Communist Bloc, the proportion of women in parliaments was very high, given the fact that these authoritarian systems were built on egalitarian ideologies. After the fall of Communism, as parliaments in post-communist countries gained real power, the percentage of female seats sharply fell. Table 6, column2, shows that our main results are preserved when excluding socialist countries (column1 reports the full sample estimated coefficients). Another concern is whether Muslim countries, where women are sometimes prevented from public activities, may affect our results. Countries with

[34]Conservative religious ideologies usually prevent women from public activities. Islamic law, for example, is typically acknowledged for its limited women's role in public; or catholicism which has been historically in opposition to women's enfranchisement, a first step in the achievement of equal political rights [Bertocchi (2011)].

[35]A country with a history of colonialism may exhibit slower incorporation of women into the political realm than countries that never were colonized [Paxton *et al.* (2006)].

[36]In order to use these additional instruments, a moment condition for the level equation, which implies that first differences of pre-determined explanatory variables are orthogonal to the country fixed effects, must be satisfied.

[37]A particular concern related to this method is the risk of instrument proliferation. Indeed, if on the one hand, the use of the entire set of instruments in a GMM context gives significant efficiency gains, on the other hand, a large collection of instruments could overfit endogenous variables as well as weaken the Hansen test of the instruments' joint validity. The instrument proliferation problem is particularly important in small samples, but unfortunately there is no formal test to detect it, even if a possible rule of thumb is to keep the number of instruments lower than or equal to the number of groups.

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Table 6. Estimations with SYS-GMM: Sample heterogeneity

	(1) Full sample Female parl. share	(2) No socialist countries Female parl. share	(3) No Muslim countries Female parl. share	(4) No SSA countries Female parl. share
Female parl. share (lagged)	0.8751***	0.9707***	0.8663***	0.8499***
	(0.096)	(0.071)	(0.104)	(0.111)
Norm of fem. parl. share (lagged)	1.728***	2.058***	1.815***	1.574***
	(0.559)	(0.422)	(0.587)	(0.553)
Democracy index	−0.0049	0.073*	−0.0575	−0.0372
	(0.054)	(0.039)	(0.065)	(0.067)
Skill ratio for females (lagged)	0.0142*	0.007	0.0142	0.0174**
	(8.4×10^{-3})	(8.6×10^{-3})	(8.6×10^{-3})	(8.2×10^{-3})
Year fixed effects	Yes	Yes	Yes	Yes
Country fixed effects	Yes	Yes	Yes	Yes
Observations	438	399	384	348
Number of countries	124	105	106	94
Number of Instruments	48	48	48	48
Arellano–Bond test for AR(1)	7.9×10^{-5}	2.2×10^{-5}	1.2×10^{-4}	5.3×10^{-4}
Arellano–Bond test for AR(2)	0.3207	0.8818	0.3281	0.2694
Hansen test *p*-value	0.2007	0.2596	0.2622	0.1478
Difference Hansen test *p*-value	0.543	0.729	0.489	0.336

*Significant at the 10% level; **5% level; ***1% level. Robust standard errors clustered by country in parentheses. SYS-GMM estimations.

at least 80% of the Muslim population are excluded from our sample.[38] Again, there is no evidence that heterogeneity plays any role in explaining our results. The last concern refers to the presence of Sub-Saharan Africa countries. Looking at the Global Gender

[38]We exclude from the sample: Algeria, Bangladesh, Egypt, Gambia, Iran (Islamic Republic of), Iraq, Jordan, Mali, Mauritania, Morocco, Niger, Pakistan, Senegal, Syrian Arab Republic, Tajikistan, Tunisia, Turkey, Yemen Arab Republic.

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Table 7. Alternative specification of the norm

	(1) Female parl. share	(2) Female parl. share
Female parl. share (lagged)	0.7316***	0.8511***
	(0.088)	(0.092)
Norm of fem. parl. share (lagged)		1.659***
		(0.508)
Norm of fem. parl. share, immigrants (lagged)	−0.7413	−0.4012
	(0.806)	(0.742)
Democracy index	7.8×10^{-5}	−0.0078
	(0.059)	(0.055)
Skill ratio for females (lagged)	0.0151**	0.0138**
	(6.5×10^{-3})	(6.9×10^{-3})
Year fixed effects	Yes	Yes
Country fixed effects	Yes	Yes
Observations	440	438
Number of countries	124	124
Number of Instruments	48	62
Arellano–Bond test for AR(1)	5.1×10^{-5}	9.2×10^{-5}
Arellano–Bond test for AR(2)	0.2683	0.3158
Hansen test p-value	0.1682	0.3459
Difference Hansen test p-value	0.773	0.967

*Significant at the 10% level; **5% level; ***1% level. Robust standard errors clustered by country in parentheses. SYS-GMM estimations.

Gap Index (2017), which considers how equitably the available income, resources, and opportunities are distributed between women and men, Sub-Saharan African countries score on average in the lower middle range of the Global Gender Gap Index, and they are characterized by a wider range of gender gap outcomes than any other region. Then, we exclude Sub-Saharan African countries. Our results are preserved.[39]

5.6 Robustness tests: alternative specification of the norm

So far, we have focused on the effects of emigrants' destination-country characteristics on female parliamentary share at the origin. A limited set of studies [e.g., Collier (2013)] emphasize the "transfer of norms" effects that immigrants can induce from origin to

[39]Sub-Saharan African countries which have been excluded: Benin, Botswana, Burundi, Cameroon, Central African Republic, Congo, Cote d'Ivoire, Democratic Republic of Congo, Gabon, Gambia, Ghana, Kenya, Lesotho, Malawi, Mali, Mauritania, Mauritius, Mozambique, Namibia, Niger, Rwanda, Senegal, South Africa, Sudan, Swaziland, Togo, Uganda, Tanzania, Zambia, Zimbabwe.

destination countries. In Table 7, we introduce in our baseline specification a term, which allows to capture the norms that immigrants bring with them when they migrate.[40] On average, the correlation between this term and the (emigration) norm is small (0.2287), so that both variables can be tested jointly. Controlling for the norm of immigrants does not affect our main results. In addition, the estimated coefficient of the norm of immigrants is negative (probably capturing the fact that over the last decades, an increase of immigrants from developing countries to OECD countries has occurred), but not statistically significant.[41]

5.7 Robustness tests: annual data

In this section, previous results are tested using annual instead of 10-year data. The latter have been chosen in the benchmark analysis for two reasons. First of all, because migration data are only available by decade. Secondly, migrants need to live in the host country for a certain period of time before assimilating and being able to transmit new values to their country. In other words, using 10-year data allows for a longer period of the occurrence of a "transfer of norm" mechanism.

However, in doing so, some important information regarding the annual political evolution by country is not taken into account. This can be particularly relevant if political legislatures last less than 5 years because of political instability or historical events such as geographical split or internal conflicts. This being said, previous estimation results have been tested using yearly data.

Some important changes were required. First of all, since migration data are available by decade, while political data contain yearly observations, the original migration decennial matrix has been extended by interpolation. Missing migration yearly data have been computed applying a constant annual rate of growth within each decade. This strategy looks globally reasonable because of two reasons. Bilateral migration stocks are very persistent which means that they vary slowly and smoothly over time. Bilateral migration stocks are used as weights in the norm whose temporal variability depends on two sources: heterogeneous bilateral migration stocks and country-specific change in the difference in the actual proportion of female parliamentary seats.[42] Secondly, also the Barro and Lee dataset, which is available every 5 years, has been interpolated using the same technique as for migration data.[43] Thirdly, migration data are complemented by the UN migration dataset for the period 2000–2010.

Table 8 shows estimation results using annual data across different econometric specifications. Columns (1) to (3) show FE results. The norm of female parliamentary share is always positive and significant. Column (4) shows SYS-GMM results using internal instruments. In SYS-GMM estimation, the lagged dependent variable is instrumented using its own sixth lag. As on average new elections occur every 5 years, the sixth lag allows to consider as instruments the parliamentary share

[40] We define a specular norm: Norm of fem. parl. share, immigrants (lagged) $= \sum_j \left[\frac{\text{mig}_{ji,t-10}}{\text{pop}_{i,t-10}} \times (\text{seats}^F_{j,t-10} - \text{seats}^F_{i,t-10}) \right]$, where mig_{ji} are migrants from country j to country i.

[41] Similarly, in Docquier *et al.* (2014), immigration does not induce significant changes in HIV prevalence rates at destination.

[42] A similar strategy has been applied by Docquier *et al.* (2014).

[43] We basically computed a five-yearly growth rate and apply it as constant to each missing human capital yearly observation.

Table 8. Estimations with annual data

	(1)	(2)	(3)	(4)
	FE	FE	FE	SYS-GMM
	Female parl. share	Female parl. share	Female parl. share	Female parl. share
Female parl. share (lagged)	0.9322***	0.9201***	0.9246***	0.9629***
	(9.8×10^{-3})	(0.013)	(0.014)	(0.021)
Norm of fem. parl. share (lagged)	0.2647**	0.4301***	0.5003***	0.3505***
	(0.111)	(0.089)	(0.14)	(0.12)
Democracy index		−0.0686***	−0.0608***	−0.0033
		(0.019)	(0.02)	(6.5×10^{-3})
Legal election		0.9568***	0.9706***	0.9923***
		(0.099)	(0.099)	(0.103)
Skill ratio for females (lagged)			0.0041**	0.0044**
			(1.6×10^{-3})	(1.8×10^{-3})
Year fixed effects	Yes	Yes	Yes	Yes
Country fixed effects	Yes	Yes	Yes	Yes
Observations	6,298	4,377	3,867	3,932
Number of countries	177	154	126	127
Number of instruments				293
Arellano–Bond test for AR(1)				1.6×10^{-8}
Arellano–Bond test for AR(2)				0.9563
Hansen test p-value				1

*Significant at the 10% level; **5% level; ***1% level. Robust standard errors clustered by country in parentheses.

of previous elections. Our variable of interest, i.e., the lagged norm of female parliamentary share, is instrumented using its own sixth lag.[44] Again, our variable of interest is positive and statistically significant with a lower coefficient with respect to the 10 years data analysis.

[44]Human capital is also treated as endogenous and instrumented using its own sixth lag, while democracy and a legal election dummy is considered as exogenous. It should be noted that the number of instruments is larger than the number of groups, and the Hansen test p-value is very high.

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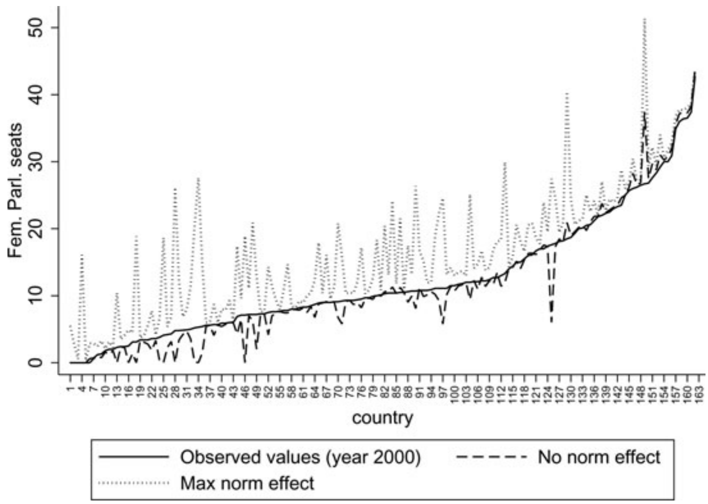

Figure 4. Predicting the female parliamentary shares using two counterfactuals.

6. Counterfactual analysis

In the main text of the paper, we found that the impact of migration in transferring political norms appears to be statistically relevant. In order to further investigate and assess the importance of this effect at the country-specific level, we simulate the counterfactual female parliamentary share obtained in two extreme cases. We consider first a simulated scenario in which migration is set equal to 0 (i.e., no transfers of norms occur). Secondly, we consider another scenario where we assume that all migrants are sent to the destination country with the highest female parliamentary share in our sample (Sweden). The latter case allows us to assess the possible maximum effect of the norm in transferring political values.

To set up the two environments, we start from our empirical model, and we consider the estimated coefficients in the baseline regression in SYS-GMM regressions (column 1 in Table 5). In particular, we focus on the short-run coefficients obtained in the estimations with data by decade and with internal instrumentation.

Let us set up the first counterfactual environment considering the baseline empirical model:

$$\text{seats}^F_{i,t} = \alpha \text{seats}^F_{i,t-10} + \beta^* \text{indexseats}^F_{i,t-10} + \sum_{i=1}^{n} \rho_i R_{i,t} + \mu_i + \varphi_t + \epsilon_{i,t} \tag{2}$$

where

$$\text{indexseats}^F_{i,t-10} = \sum\nolimits_j \left[\frac{\text{mig}_{ij,t-10}}{\text{pop}_{i,t-10}} \times (\text{seats}^F_{j,t-10} - \text{seats}^F_{i,t-10}) \right]$$

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Assuming no migration, we have:

$$\text{seats}_{i,t}^{F} = \alpha \text{seats}_{i,t-10}^{F} + \beta^{*}\text{indexseats}_{i,t-10}^{F} + \sum_{i=1}^{n} \rho_i R_{i,t} + \mu_i + \varphi_t + \epsilon_{i,t} \tag{3}$$

where $\text{indexseats}_{i,t-10}^{F} = 0$. Taking the difference between (2) and (3) gives us the change in the female parliamentary seats:

$$\Delta \text{seats}_{i,t}^{F} \equiv \text{seats}_{i,t}^{F} - \text{seats}_{i,t}^{F} = -\beta^{*}\text{indexseats}_{i,t-10}^{F} \tag{4}$$

which can be re-written as:

$$\text{seats}_{i,t}^{F} = \text{seats}_{i,t}^{F} - \beta * \text{indexseats}_{i,t-10}^{F} \tag{5}$$

For simplicity, we consider data for the year 2000 and $\beta = 1.728$ and we construct the counterfactual values for female parliamentary seats ($\text{seats}_{i,t}^{F}$) in each country in the case of no migration (i.e., no transfers of norms). The dashed line in Figure 4 shows the counterfactual value for female parliamentary seats in the case of no migration.[45] As we can see from the graph, for some countries, the counterfactual value for female parliamentary participation in 2000 is lower than the observed value in the same year of interest. This is especially true for countries with lower female political empowerment, in particular developing countries, for which migration is shown to be particularly relevant in improving women's conditions. For other countries with a high share of females in parliament, such as Sweden, Denmark, Finland, Norway, the Netherlands, however, the counterfactual value is higher than the observed value, meaning that emigration may in some cases decrease female political empowerment. On average, with respect to the median of the distribution, female parliamentary seats will decrease from about 10% to 9.38%.[46,47]

In a second and symmetric counterfactual experiment, we compute the maximum effect the transfer of norm mechanism could have. We assume that all migrants are sent to the country with the most feminized parliament (i.e., all migrants are sent to Sweden which has 42.7% of female parliamentary seats in 2000).[48] As the dotted line

[45]Table E1 of Appendix E shows the counterfactual results for all the countries with no missing data of female parliamentary seats in 2000 and that belong to our estimated sample. See the table for the correspondence between country ranking and the name of the country.

[46]To give some numerical examples, the countries that "lose" the most without migration are developing countries. For example, Albania would decrease from 5.2 to 0.14. For Lesotho would decrease from 3.8 to 0.27. For Turkey, however, female parliamentary seats would decrease from 4.2 to 2.24. For countries such as Namibia, there are almost no changes (from 25 to 24.99). On the other hand, among the countries which would gain more from a "no transfer of norm" environment, we find Sweden, with an increase in female parliamentary seats from 42.7 to 43.74; Norway (from 36.4 to 37.19); but also Latvia (from 17 to 17.58); New Zealand (from 29.2 to 31); Guyana (from 18.5 to 20.92); Grenada (from 26.7% to 37.43%). Table E1 of Appendix E shows the counterfactual results for all the countries.

[47]If we perform a similar counterfactual using the estimated long-run coefficient $\beta = 13.8$, we obtain that on average, with respect to the median of the distribution, the share of female parliamentary seats will decrease from about 10% to 6.91%.

[48]In this case, equation (5) becomes: $\text{seats}_{i,t}^{F} = \text{seats}_{i,t}^{F} + \beta * (\text{indexseats}_{iswe,t-10}^{F} - \text{indexseats}_{i,t-10}^{F})$, where $\text{indexseats}_{iswe,t-10}^{F} = [(\text{mig}_{iswe,t-10}/\text{pop}_{i,t-10}) \times (\text{seats}_{swe,t-10}^{F} - \text{seats}_{i,t-10}^{F})]$. For simplicity, also in this case, we consider data for the year 2000 and $\beta = 1.728$.

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in Figure 4 shows, all the countries have a higher share of female parliamentary seats. On average, with respect to the median of the distribution, female parliamentary seats will increase from about 10% to 14.7%.[49,50]

7. Conclusion

Women make up more than half of the population in the world. Female electorates have globally grown up in the last two decades but yet continue to be under-represented in political decision-making bodies at all levels. The World Development Report 2012 states that gender equality matters for development-enhancing productivity, creating a better environment for the next generation and making institutions more representative. In addition, there is evidence [Clots-Figueras (2011), Thomas (1991)] that women in politics improve development outcomes for women themselves, children, and families.

The World Bank (2011) wonders whether "globalization can help" in fostering gender equality. In this paper, we have partly answered this by providing some evidence on how a globalized outcome such as international migration has contributed to the increase of female parliamentary participation from 1970 to 2010. In other words, international migrants have acted as "informational" channels able to transfer foreign values, create favorable opportunities, reshape attitudes, and create new norms about women in the origin country.

Following the brand new strand of literature on "transfers of norms" [e.g., Spilimbergo (2009), Beine *et al.* (2013)], we have applied the same mechanism to female political participation. To this end, we estimated a dynamic model by decades in which female access to Parliament depends on traditional covariates plus international migration. The empirics contains two important insights. First of all, the norm (through which foreign female parliamentary participation is propagated at origin) has been constructed in such a way that the origin country takes advantage of the political environment at the destination just if the female political conditions at destination are better than those at the origin. Secondly, endogenous issues are taken into account.

Results, which are robust to different geopolitical specifications, show that female political empowerment can be accounted as another migration non-economic externality, suggesting that the launch of domestic public actions can also be supported by the role of active national people from abroad.

Finally, our study can be seen as a complement to the ones which examine the effect of emigration on source-country fertility [e.g., Beine *et al.* (2013)]: migration-induced women empowerment could be seen as a specific channel through which international migration conveys a transfer of fertility norms across countries. Empirical evidence indicates that female legislators spend more in health and

[49]Remarkably effects would occur for Armenia (from 3.1 to 18.837) and Morocco (from 0.6 to 2.90); and the minimum effect pertains to China (from 21.08 to 21.93). Table E1 of Appendix E shows the counterfactual results for all the countries.

[50]If we perform a similar counterfactual using the estimated long-run coefficient $\beta = 13.8$, we obtain that on average, with respect to the median of the distribution, the share of female parliamentary seats will increase from about 10% to 32.9%. However, the long-run counterfactual provides unreliable results for countries which belong to the top 10% of the female parliamentary share distribution, with simulated values about 100%.

education, and favor "women-friendly" laws [Clots-Figueras (2011)]. Favoring equal gender rights can affect women's bargaining position within their households, and thus lead to lower fertility [Doepke *et al.* (2012)].

Acknowledgements. An early version of this paper circulated as IRES Discussion Papers 2012001, Universite catholique de Louvain [Lodigiani and Salomone (2012)]. This work has been supported by the European Social Fund ("Dote ricerca") from the Lombardia region and the Belgian French-speaking Community (convention ARC 09/14-019 on "Geographical Mobility of Factors"). We thank Massimiliano Bratti, Frédéric Docquier, Giovanni Facchini, Emanuele Forlani, Tommaso Frattini, Julien Martin, Elie Murard, Matteo Picchio, and all the participants to the Conference on "The Economics and Politics of Immigration", the 2nd TEMPO Conference on International Migration, the ARC Workshop on "Geographical Mobility of Workers and Firms", the seminar held in the Department of Economics at the University of Modena and Reggio Emilia and the LdA/Paolo Baffi research seminar at Bocconi University, for their constructive comments. We are also highly thankful to Chris Parsons and Pamela Paxton for having shared their data with us.

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Appendix A

List of countries

The geographical sample of interest as in the first two columns of Table 1 in the main text is equal to 169 countries. The 169 countries are: Albania, Algeria, Angola, Antigua and Barbuda, Argentina, Armenia, Australia, Austria, Azerbaijan, Bahamas, Bangladesh, Barbados, Belarus, Belgium, Belize, Benin, Bhutan, Bolivia, Bosnia and Herzegovina, Botswana, Brazil, Bulgaria, Burkina Faso, Burundi, Cambodia, Cameroon, Canada, Cape Verde, Central African Republic, Chad, Chile, China, Colombia, Comoros, Congo, Costa Rica, Cote d'Ivoire, Croatia, Cuba, Cyprus, Czech Republic, Czechoslovakia, Democratic People's Republic of Korea, Democratic Republic of the Congo, Denmark, Djibouti, Dominican Republic, Ecuador, Egypt, El Salvador, Equatorial Guinea, Eritrea, Estonia, Ethiopia, Fiji, Finland, France, Gabon, Gambia, Georgia, Germany, Ghana, Greece, Grenada, Guatemala, Guinea-Bissau, Guyana, Haiti, Honduras, Hungary, Iceland, India, Indonesia, Iran (Islamic Republic of), Iraq, Ireland, Israel, Italy, Jamaica, Japan, Jordan, Kazakhstan, Kenya, Kiribati, Kyrgyzstan, Lao People's Democratic Republic, Latvia, Lebanon, Lesotho, Lithuania, Luxembourg, Madagascar, Malawi, Malaysia, Maldives, Mali, Malta, Mauritania, Mauritius, Mexico, Micronesia (Federated States of), Mongolia, Morocco, Mozambique, Namibia, Nepal, The Netherlands, New Zealand, Nicaragua, Niger, Nigeria, Norway, Pakistan, Panama, Papua New Guinea, Paraguay, Peru, Philippines, Poland, Portugal, Republic of Korea, Republic of Moldova, Romania, Russian Federation, Rwanda, Saint Lucia, Saint Vincent and the Grenadines, Samoa, Sao Tome and Principe, Senegal, Serbia and Montenegro, Seychelles, Singapore, Slovakia, Slovenia, Socialist Federal Republic of Yugoslavia, Solomon Islands, South Africa, Spain, Sri Lanka, Sudan, Suriname, Swaziland, Sweden, Switzerland, Syrian Arab Republic, Tajikistan, Thailand, The former Yugoslav Republic of Macedonia, Togo, Tonga, Trinidad and Tobago, Tunisia, Turkey, Turkmenistan, USSR, Uganda, Ukraine, United Kingdom of Great Britain and Northern Ireland, United Republic of Tanzania, United States of America, Uruguay, Uzbekistan, Vanuatu, Venezuela, Viet Nam, Yemen Arab Republic, Zambia, Zimbabwe.

Appendix B

Construction of the norms

In order to construct the "norms" in equation (1) in the main text, the final matrix should be perfectly balanced. By final matrix we mean the matrix comprising: migration data [from Ozden *et al.* (2011) and UN data for year 2010], political data [from Paxton *et al.* (2006) and World Development Indicators from 2003 to 2010], total population (from World Population Prospects: *The 2012 Revision* United Nations 2013).

For migration data, data for Czechoslovakia, the Socialist Federal Republic of Yugoslavia, and USSR were not available in the original dataset. We reconstructed missing observations aggregating migration data for the countries belonging to them before the political scission. So for Czechoslovakia before 1993 (replaced by missing values then), we aggregated data from Czech Republic and Slovakia. For the Socialist Federal Republic of Yugoslavia before 1992 (replaced by missing values then), we aggregated data from Bosnia and Herzegovina, Croatia, Serbia and Montenegro, Slovenia, and the Former Yugoslav Republic of Macedonia. For USSR before 1991 (replaced by missing values then), we aggregated data from Ukraine, Russian Federation, Uzbekistan, Kazakhstan, Belarus, Azerbaijan, Georgia, Tajikistan, Republic of Moldova, Kyrgyzstan, Lithuania, Turkmenistan, Armenia, Latvia, and Estonia. For political data, we explained in Appendix D how we deal with "true missing values", while for the other missing cells (about 31% of the sample) due to political instability, coup d'etat, dictatorship war, presence of "false elections", lack of sovereignty due to colonialism, we have kept as missing. Once we have dealt with missing values in each dataset to make each of them balanced, we have merged the three of them to obtain the final dataset. Then the norms have been constructed.

Appendix C

Difference with previous studies

Our empirical specification mainly refers to the work by Spilimbergo (2009). Following step by step his dynamic empirical specification, equation (1) in the main text becomes:

$$\begin{aligned} \text{seats}^F_{i,t} &= \alpha \text{seats}^F_{i,t-10} + \beta \text{indexseats}^F_{i,t-10} + \gamma \text{migrate}_{i,t-10} \\ &+ \eta \text{inter} + \sum_{i=1}^{n} \rho_i R_{i,t} + \mu_i + \epsilon_{i,t} \end{aligned} \tag{C.1}$$

where:

- $\text{indexseats}^F_{i,t-10} = \sum_j \left[\left(\text{mig}_{ij,t-10} / \sum_i \text{mig}_{ij,t-10} \right) \times (\text{seats}^F_{j,t-10} - \text{seats}^F_{i,t-10}) \right]$. In Spilimbergo (2009), the exact corresponding variable $\text{indexseats}^F_{i,t-10}$ would be constructed as the weighted average of the female parliamentary share in the host countries, e.g., $\text{indexseats}^F_{i,t-10} = \sum_j \left[\left(\text{mig}_{ij,t-10} / \sum_i \text{mig}_{ij,t-10} \right) \times \text{seats}^F_{i,t-10} \right]$. Just considering the level of female parliamentary share at destination would have prevented account being taken of political asymmetries between origin and destination countries. With this specification, indeed, the "transfer of norm" is always positive if $\text{seats}^F_{i,t-10}$ is >0, apart from the level of female parliamentary share at origin. In considering the difference between female parliamentary shares between countries of destination and origin, however, we assume that there is a "positive transfer" only when migrants reside in countries where female political conditions are better, and that the transfer is higher, the greater the political difference between the two countries.
- $\text{migrate}_{i,t-10}$ is the ratio between $\left(\sum_i \text{mig}_{ij,t-10} / \text{pop}_{i,t-10} \right)$
- inter is the interaction term and corresponds to $(\text{migrate}_{i,t-10} \times \text{indexseats}^F_{i,t-10})$

Equation (1) in the main text and equation (C.1) are symmetric. The only difference consists in the construction of the "norm". In other words, in equation (C.1), the average female parliamentary share at the destination is computed as:

$$\text{indexseats}^F_{i,t-10} = \sum_j \left[\frac{\text{mig}_{ij,t-10}}{\sum_i \text{mig}_{ij,t-10}} \times (\text{seats}^F_{j,t-10} - \text{seats}^F_{i,t-10}) \right] \tag{C.2}$$

where the weights are given by emigration shares. In equation (1), however, the weights are given by emigration rates. In other terms, we substitute $\text{mig}_{ij,t-10} / \sum^i \text{mig}_{ij,t-10}$ with $\text{mig}_{ij,t-10}/\text{pop}_{i,t-10}$. So the norm in the benchmark specification becomes as follows:

$$\text{indexseats}^F_{i,t-10} = \sum_j \left[\frac{\text{mig}_{ij,t-10}}{\text{pop}_{i,t-10}} \times (\text{seats}^F_{j,t-10} - \text{seats}^F_{i,t-10}) \right] \tag{C.3}$$

Obviously, due to the different nature of Spilimbergo's norm, equation (C.1) also contains the total migration rate calculated as the ratio between total aggregate migration from country i over total population in country i and the interaction term between the total migration rate and the average index of female political participation at the destination.

As already explained in the text, the lagged index of female parliamentary share à la Spilimbergo affects the female parliamentary participation at time t but the interaction with migration rate is not significant as in Beine *et al.* (2013) and sometimes in Spilimbergo (2009). The lack of significance can be due to collinearity since migration rate appears three times as regressor: alone, then it is used as a weight in the norm, and as a multiplicative term (as migration share) in the interaction term. Indeed, the collinearity between the interaction term and the migration rate is more than 90%. A way to avoid collinearity is an alternative construction of the norm as in equation (1) of the main text.

Appendix D

Addressing the "true missing" values in Paxton et al. (2006)

Here is the list of countries which contains true missing values:

- Belarus: data from 1995 to 1999 are true missing. The missing cells have been complemented by the CPDS II [Armingeon and Careja (2008)], which covers 28 post-communist countries for the 1989–2008 period.
- Benin: years 1960, 1961, 1962, 1964 are true missing. Following Benin political and historical information, we transformed true missing values into missing.
- Bhutan: year 1996 is a true missing. It has been replaced with observed data from IPU (01/1996 elections).
- Bosnia-Herzegovina: true missing for 1996–1999. We replaced years 1996–1997 with missing data in accordance with the female parliamentary share in contiguous countries, e.g., Croatia, Serbia, and Montenegro. While years 1998–1999 have been complemented by the CPDS II [Armingeon and Careja (2008)], which covers 28 post-communist countries for the 1989–2008 period.
- Cambodia: year 2003 is a true missing. It has been replaced by the observed positive value from IPU, election 07-2003.
- The Democratic Republic of the Congo: true missing data for 1992–1993, 2000–2002, and 2003. The average female parliamentary share between years 1991 and 1994 replaces the first missing biennium since data are quite stable. True missing data from 2000 to 2002 have been transformed into missing because of political instability and civil war. True missing for 2003 has been replaced by data from IPU election 08/2003.
- Gambia: data from 1977 to 1981 are true missing. They have been replaced by 0 as women started to stand for election from 1982.
- Guinea: true missing data from 1981 to 1984 have been transformed into missing for political instability.
- Guyana: true missing data from 1964 to 1967 have been changed into missing because the first parliamentary election occurs in 1968 [from Golder (2005)].
- Kiribati: true missing for year 2003 has been substituted with the observed positive value from IPU, election 05/2003.
- Latvia: true missing values from 1990 to 1992 have been replaced by missing. Latvia gained independence in 1991. According to Golder (2005) and to the Comparative Political Data Set II from Armingeon and Careja (2008), 1993 is the year of the first election after the fall of communist rule.
- Liberia: the country has been dropped because true missing cells exceed 50%.
- Mali: true missing values from 1988 to 1990 have been replaced by missing because of political instability due to the dictatorship of Moussa Traori, before a coup d'etat.
- Marshall Islands: true missing data from 1995 to 1998 have been replaced by values of the previous (1994) and following (1999) elections which are equal.
- MyanMar: true missing data from 1960 to 1963 have been replaced by missing because of political instability. True missing data for 1985–1987 have been complemented by the average value from previous years.
- Nauru: true missing data from 1992 to 1994 have been replaced by the mean between previous and following years. While true missing for 2003 has been replaced by IPU value for election 05/2003.
- Niger: true missing value for 1992 has been replaced by missing because of political instability.
- Nigeria: all the true missing values have been replaced by missing because of strong political instability and civil wars.
- Pakistan: true missing values from 1960 to 1972 have been replaced by missing because of political instability and lack of constitution. While the true missing for year 1996 has been replaced by the 1995's value, (stable data in the 90s).
- Peru: the true missing value for year 2000 has been replaced by the observed data from IPU (election 04-2000).

- Rwanda: true missing data from 1966 to 1971 have been replaced by missing because of political instability.
- Seychelles: true missing data for the biennium 1991–1992 have been replaced by the value for 1990, for historical reasons. In 1977, a coup d'etat ousted the first president of the republic, James Mancham, who was replaced by France Albert René. The 1979 constitution declared a socialist one-party state, which lasted until 1991. The first draft of a new constitution failed to receive the requisite 60% of voters in 1992, but an amended version was approved in 1993.
- Sri Lanka: true missing data for year 2000 have been replaced by the average between 1999 and 2001 values.
- Tonga: true missing data for 2002–2003 have been replaced by the observed value of the election 03/2002 from IPU.
- Uganda: true missing values from 1962 to 1965 have been replaced by missing because of political instability.
- Tanzania: true missing values from 1965 to 1969 have been replaced by missing because of political instability.
- Vanuatu: true missing values for 1995–1997 have been replaced by missing because of political instability.

Appendix E

Country rankings

Table E1. Counterfactual analysis

Country	Country ranking	Baseline	No migration	Migration to Sweden
Jordan	1	0	0.0	5.5
Micronesia (Fed. States of)	2	0	0.0	2.7
Vanuatu	3	0	0.0	0.5
Tonga	4	0	0.0	16.1
Djibouti	5	0	0.0	0.3
Morocco	6	0.6	0.0	2.9
Yemen Arab Republic	7	0.7	0.6	2.9
Niger	8	1.2	1.0	2.5
Mauritania	9	1.3	0.7	3.3
Papua New Guinea	10	1.8	1.6	2.1
Gambia	11	2	1.2	3.1
Solomon Islands	12	2	1.8	2.3
Lebanon	13	2.3	0.0	10.5
Chad	14	2.4	2.3	3.5
Egypt	15	2.4	2.3	4.1
Paraguay	16	2.5	0.0	4.9

(*Continued*)

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Table E1. (*Continued.*)

Country	Country ranking	Baseline	No migration	Migration to Sweden
Swaziland	17	3.1	1.4	4.5
Armenia	18	3.1	0.0	18.8
Nigeria	19	3.4	3.3	3.7
Iran (Islamic Rep. of)	20	3.4	3.0	4.1
Algeria	21	3.4	2.8	5.6
Haiti	22	3.6	1.8	7.8
Kenya	23	3.6	3.3	4.1
Lesotho	24	3.8	0.3	6.6
Samoa	25	4.1	0.0	18.7
Turkey	26	4.2	2.2	5.4
Singapore	27	4.3	3.2	6.9
Saint Vincent and the Grenadines	28	4.8	0.0	26.2
Kiribati	29	4.8	3.4	12.2
Sri Lanka	30	4.85	4.2	6.9
Togo	31	4.9	4.7	7.9
Equatorial Guinea	32	5	3.5	11.4
Albania	33	5.2	0.1	20.2
Antigua and Barbuda	34	5.3	0.0	27.6
Cyprus	35	5.4	1.5	14.7
Thailand	36	5.6	5.5	6.0
Cameroon	37	5.6	5.5	6.1
Mauritius	38	5.7	4.1	8.7
Brazil	39	5.7	5.6	6.0
Republic of Korea	40	5.9	5.4	8.0
Nepal	41	5.9	5.7	7.8
Benin	42	6	6.1	9.3
Maldives	43	6	6.0	6.1
Belize	44	6.9	4.3	17.4
Guatemala	45	7.1	6.5	9.6
Bosnia and Herzegovina	46	7.1	0.0	18.9
Uzbekistan	47	7.2	7.0	11.1
Georgia	48	7.2	6.2	20.9
Serbia and Montenegro	49	7.2	2.0	12.8

(*Continued*)

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Table E1. (*Continued.*)

Country	Country ranking	Baseline	No migration	Migration to Sweden
Japan	50	7.3	7.2	7.6
Central African Republic	51	7.3	7.3	7.7
Macedonia	52	7.5	4.2	14.3
Russian Federation	53	7.6	7.1	11.4
Iraq	54	7.6	7.2	9.4
Ethiopia	55	7.7	7.7	7.9
Guinea-Bissau	56	7.8	7.5	10.9
Ukraine	57	7.8	7.4	14.7
Indonesia	58	8	7.9	8.4
Madagascar	59	8	7.9	8.2
Cambodia	60	8.2	7.8	9.2
Gabon	61	8.3	8.2	8.9
Hungary	62	8.3	7.5	10.0
Cote d'Ivoire	63	8.5	8.4	10.4
Greece	64	8.7	6.8	12.3
Republic of Moldova	65	8.9	8.8	17.9
Ghana	66	9	8.9	10.4
Burkina Faso	67	9	9.1	16.1
India	68	9	9.0	9.5
Bangladesh	69	9.1	9.1	11.2
Sao Tome and Principe	70	9.1	6.7	20.8
Malta	71	9.2	5.8	15.9
Malawi	72	9.3	9.1	10.3
Bhutan	73	9.3	9.3	10.5
Zimbabwe	74	9.3	8.7	10.4
Honduras	75	9.4	9.0	12.1
El Salvador	76	9.5	8.2	17.2
Venezuela	77	9.7	9.5	10.3
Sudan	78	9.7	9.7	10.8
Panama	79	9.9	9.4	12.7
Kyrgyzstan	80	10	10.4	18.3
Zambia	81	10.1	9.9	10.8
Belarus	82	10.3	10.6	20.4

(*Continued*)

Table E1. (*Continued.*)

Country	Country ranking	Baseline	No migration	Migration to Sweden
Malaysia	83	10.4	10.6	13.1
Kazakhstan	84	10.4	11.3	24.2
Syrian Arab Republic	85	10.4	10.3	11.8
Azerbaijan	86	10.5	11.2	21.5
Mongolia	87	10.5	10.5	10.6
Lithuania	88	10.6	9.0	17.5
Romania	89	10.7	10.0	13.2
Barbados	90	10.7	8.1	26.3
Bulgaria	91	10.8	11.2	16.6
Nicaragua	92	10.8	9.9	15.3
Chile	93	10.8	10.1	12.0
France	94	10.9	10.6	12.0
Saint Lucia	95	11.1	10.2	18.9
Trinidad and Tobago	96	11.1	9.3	22.3
Cape Verde	97	11.1	5.8	24.6
Italy	98	11.1	10.1	13.1
Tunisia	99	11.5	11.4	14.2
Bolivia	100	11.5	10.7	13.0
Israel	101	11.7	11.5	13.5
Colombia	102	11.8	11.6	13.7
Congo	103	12	11.9	13.0
Ireland	104	12	9.5	25.1
Senegal	105	12.1	12.3	14.1
Uruguay	106	12.1	11.0	14.7
Mali	107	12.2	12.8	16.7
Slovenia	108	12.2	11.4	13.8
Philippines	109	12.4	12.3	14.2
Slovakia	110	12.7	11.7	16.9
Tajikistan	111	12.7	13.4	18.1
Poland	112	13	11.4	18.3
Jamaica	113	13.3	12.0	29.9
United States of America	114	14	14.0	14.3
Burundi	115	14.4	14.2	15.8

(*Continued*)

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Table E1. (*Continued.*)

Country	Country ranking	Baseline	No migration	Migration to Sweden
Bahamas	116	15	15.2	20.6
Czech Republic	117	15	14.0	18.1
Angola	118	15.5	15.3	16.6
Mexico	119	16	16.3	20.6
Dominican Republic	120	16.1	16.3	20.8
Luxembourg	121	16.7	16.1	18.6
Botswana	122	17	16.8	17.8
Latvia	123	17	17.6	23.9
Ecuador	124	17.4	17.3	19.5
Suriname	125	17.6	6.0	27.4
Estonia	126	17.8	17.3	24.6
Uganda	127	18.1	18.5	19.4
United Kingdom	128	18.2	18.0	20.9
Guyana	129	18.5	20.9	40.4
Portugal	130	18.7	19.0	24.3
Costa Rica	131	19.3	19.5	20.6
Peru	132	20	20.2	21.2
Democratic People's Republic of Korea	133	20.1	20.6	21.4
Croatia	134	20.5	19.7	25.1
Canada	135	20.6	21.0	22.5
Lao People's Dem. Republic	136	21.2	21.9	24.2
China	137	21.8	21.8	21.9
Eritrea	138	22	23.8	27.0
United Republic of Tanzania	139	22.3	22.4	22.7
Switzerland	140	22.5	22.8	24.4
Australia	141	23	23.2	24.0
Belgium	142	23.3	23.5	24.6
Seychelles	143	23.5	24.6	28.8
Namibia	144	25	25.0	26.0
Rwanda	145	25.7	26.2	27.0
Turkmenistan	146	26	28.3	30.4

(*Continued*)

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Table E1. (*Continued.*)

Country	Country ranking	Baseline	No migration	Migration to Sweden
Viet Nam	147	26.2	26.6	27.2
Argentina	148	26.5	26.8	27.3
Grenada	149	26.7	37.4	51.4
Austria	150	26.8	27.5	29.3
Cuba	151	27.6	29.7	32.1
Spain	152	28.3	28.8	29.5
New Zealand	153	29.2	31.0	34.0
South Africa	154	30	30.2	30.6
Mozambique	155	30	30.8	31.7
Germany	156	30.9	32.0	32.9
Iceland	157	34.9	35.5	36.8
The Netherlands	158	36	37.2	37.8
Norway	159	36.4	37.2	37.7
Finland	160	36.5	37.3	38.0
Denmark	161	37.4	38.3	38.6
Sweden	162	42.7	43.7	43.7

Cite this article: Lodigiani E, Salomone S (2020). Migration-induced transfers of norms: the case of female political empowerment. *Journal of Demographic Economics* **86**, 435–477. https://doi.org/10.1017/dem.2020.7